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Clustering of Atherogenic Behaviors in Coffee Drinkers

Elizabeth M. Puccio, MD, Janice B. McPhillips, MS, Elizabeth Barrett-Connor, MD,
and Theodore G. Ganiats, MD

Abstract: We studied the clustering of coffee consumption and selected atherogenic behaviors in older adults living in a southern California community. Men were somewhat more likely to drink caffeinated coffee while women were more likely to drink decaffeinated coffee. In men, but not women, caffeinated coffee drinking decreased with age and decaffeinated coffee drinking increased. Caffeinated coffee drinkers drank more alcohol, consumed more dietary saturated fats and cholesterol, were more likely to be current smokers and less likely to be current exercisers than were non-coffee drinkers. Smoking and exercise also showed a dose-response relationship to the amount of caffeinated coffee consumed. Risk factor levels among drinkers of decaffeinated coffee were more like those of caffeinated coffee than non-drinkers. These data illustrate the clustering of atherogenic behaviors with coffee drinking and highlight their potential importance in interpreting the growing body of literature about coffee and health. (*Am J Public Health* 1990; 80:1310–1313).

Introduction

Whether coffee drinking is a risk factor for coronary heart disease continues to be controversial. Some epidemiologic evidence has linked heavy coffee consumption to elevations in total serum cholesterol or low density lipoprotein cholesterol,[1–3] suggesting a mechanism for the association of coffee and coronary heart disease. Several investigators have reported an increased risk of coronary heart disease morbidity and mortality with heavy coffee intake,[4–8] but others have questioned whether this association is independent of cigarette smoking.[9–12] Only two previous studies[13,14] have examined coffee use in relation to multiple behaviors other than smoking, that could themselves be associated with hyperlipidemia and/or coronary heart disease. In this paper, the coffee drinking habits of a southern California population-based cohort of older men and women were examined in relation to multiple putative heart disease risk factors to determine the extent to which coffee drinking is associated with other atherogenic behaviors.

Methods

Participants were a part of the Rancho Bernardo Heart and Chronic Disease Study, a population-based study designed to examine lifestyle and chronic disease patterns in older adults. All surviving members of the cohort who were between 40 and 84 years of age when first seen in 1972-74 were invited to participate in 1984-87, and 80 percent did so. Height and weight were measured with participants in light clothing without shoes; obesity was estimated using body mass index (kg/m^2). Data on current exercise habits, current and past smoking history, current alcohol consumption, hours of nightly sleep, and use of non-contraceptive estrogens (women only) were determined by a standardized interview administered by trained personnel. Coffee intake and diet data were ascertained as part of a self-administered food frequency questionnaire. Participants were asked to estimate their average consumption of decaffeinated and regular coffee (in cups per day) and their usual intake (as an adult over 45 years of age) of 128 food items.

A crude index of dietary saturated fat and cholesterol intake was made based upon reported consumption of foods selected by the American Heart Association as food to limit in a fat-controlled, low-cholesterol diet.[15] Six responses for frequency of intake ranging from "never" to "daily" were transformed to ranks from "1" to "6" with "1" corresponding to "never" and "6" corresponding to "daily." The cholesterol and saturated fat index was created by computing an integer sum of the participants' coded estimates of consumption for the following 12 foods or food categories: beef, ham, hot dogs, eggs, milk, organ meats, cream, cheese, ice cream, butter, lard, dessert cakes.

Participants were grouped into four coffee drinking categories: those who drank no coffee, drinkers of caffeinated coffee only, drinkers of decaffeinated coffee only, or those who drank both caffeinated and decaffeinated coffee. Participants were also grouped according to whether or not they reported exercising at least three times a week, whether they currently smoked cigarettes, quit smoking, or never smoked, and for female participants, whether or not they reported having taken any oral estrogen within the past two weeks. The total number of bottles or cans of beer, glasses of wine, highballs, cocktails, mixed drinks, liqueurs or other alcoholic drinks consumed was used as an index of alcohol consumption; daily alcohol intake estimated from the past week or from usual intake from the diet questionnaire were nearly identical and only the former is shown. Reported hours of usual nightly sleep were estimated to the nearest hour.

Results are presented for 2,304 men and women ages 50-89. One-way analysis of covariance was used to compare mean age-adjusted risk factor levels among those who drank caffeinated coffee only, decaffeinated coffee only, or both caffeinated and decaffeinated coffee to those who drank no coffee. In addition, mean risk factors for caffeinated coffee drinkers were compared to decaffeinated coffee drinkers. Age-adjusted rates of categorical risk factors were computed by the direct method and differences in age-adjusted rates were tested using the Mantel-Haenszel summary chi-square with one degree of freedom. No adjustments were made to account for the effects of multiple comparisons.

Coffee consumption in those who drank caffeinated coffee only (464 men, 514 women) was divided into three categories as follows: one to two cups per day, three to four cups per day, and

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From the Department of Community and Family Medicine, Division of Epidemiology (Puccio, McPhillips, Barrett-Connor) and the Division of Family Medicine (Ganiats), University of California School of Medicine, San Diego. Address reprint requests to Elizabeth Barrett-Connor, MD, Professor and Chair, Department of Community and Family Medicine, Division of Epidemiology M-007, University of California/San Diego, La Jolla, CA 92093-0607. This paper, submitted to the Journal August 4, 1989, was revised and accepted for publication June 5, 1990.

five or more cups per day. Trends in age-adjusted behavioral risk factor with increasing coffee consumption were tested by analysis of covariance for continuous variables and by the Mantel extension test[16] for categorical variables. Because there were very few heavy drinkers of decaffeinated coffee only, analysis of trends could not be performed.

Results

As shown in Table 1, the 1,028 men and 1,276 women ages 50-89 years were similar with regard to the proportion who drank no coffee (11 percent) or caffeinated and decaffeinated coffee combined (28 percent). Men were somewhat more likely to drink caffeinated coffee only and women to drink decaffeinated coffee only, although both sex differences were small. The proportion of men who drank only decaffeinated coffee increased with age (p for trend <0.001) while the proportion who drank only caffeinated decreased (p for trend = 0.002); similar trends were not observed in women.

Mean age and sex-adjusted behaviorial risk factors for coronary heart disease (CHD) by coffee consumption category are summarized in Table 2. In both men and women, caffeinated coffee drinkers were significantly more likely to be current cigarette smokers and to consume more alcohol and saturated fats and cholesterol than non-coffee drinkers. The percent of women who exercised at least three times per week was lower in all types of coffee drinkers as compared to non-coffee drinkers; the reverse was true for men (data not shown). Neither difference was statistically significant. BMI and hours of nightly sleep did not differ significantly in coffee drinkers as compared to non-drinkers in either sex.

Age-adjusted levels of many risk factors among decaffeinated coffee drinkers were more similar to those who drank caffeinated coffee than to those who drank no coffee (Table 2). However, dietary consumption of saturated fats and cholesterol was significantly higher in caffeinated coffee drinkers as compared to decaffeinated coffee drinkers in both men and women. Women who drank caffeinated coffee had a significantly lower use of replacement estrogen than decaffeinated coffee drinkers (data not shown). Drinkers of caffeinated coffee consumed more alcohol and were more likely to be current cigarette smokers than those who drank decaffeinated coffee.

Among those who drank caffeinated coffee only, the proportion of men and women who currently smoke cigarettes increased with increasing caffeinated coffee consumption (p for trend in age-adjusted rates <0.01 for men, <0.001 for women) (Table 3). The frequency of current exercise decreased with increasing coffee consumption in both men and women although the trend was significant only among men. There were no other clear associations between the amount of coffee consumed and alcohol, dietary cholesterol, hours of nightly sleep, or post-menopausal estrogen use (among women only). The range of decaffeinated coffee drinking (one to two cups) was too narrow to be considered in a similar analysis.

Discussion

Coffee drinking may be a marker for an atherogenic diet, cigarette smoking, or sedentary lifestyle, rather than an etiologic factor in heart disease outcomes. For example, Hennekens, *et al.*,[10] reported a positive association between coffee drinking and mortality from coronary heart disease, independent of cigarette smoking, that disappeared with adjustment for daily alcohol consumption, level of physical activity, and use of coffee additives.

In the present study, there was an increased occurrence of three habits in caffeinated coffee drinkers: use of cigarettes, excess consumption of alcohol, and an atherogenic diet. There was a dose-response relationship in the extent to which some of these behaviors were practiced among drinkers of caffeinated coffee.

These results support the findings of Jacobsen and Thelle,[13] Haffner, *et al.*,[14] and others[17],[*] who found that coffee drinkers consume a more atherogenic diet than their non-coffee drinking counterparts. Carmody, *et al.*,[18] also reported that smokers and ex-smokers were more likely than non-smokers to drink greater amounts of alcohol and coffee; most individuals who reported drinking more alcohol also reported drinking more coffee, regardless of smoking habit. In a study of an industrial population, Phillips, *et al.*,[19] noted that smokers drank an average of 1.5 to 2 more cups of coffee per day and drank more alcohol weekly than non-smokers.

Others have reported that coffee drinkers have a decreased body mass relative to non-coffee drinkers.[20] In this cohort, body mass index did not differ by coffee intake, but those who drank five or more cups of caffeinated coffee per day exercised significantly less than those who drank less. Leon, *et al.*,[21] found that the use of caffeine-containing

TABLE 1—Distribution of Coffee Consumption According to Age and Sex, Rancho Bernardo, 1984–87

		No Coffee		Decaffeinated Only		Caffeinated Only		Both	
Sex, Age (years)	N	n	%	n	%	n	%	n	%
Men									
50–59	166	27	16.3	16	9.6	83	50.0	40	24.1
60–69	254	30	11.8	31	12.2	129	50.8	64	25.2
70–79	404	30	7.4	63	15.6	175	43.3	136	33.7
80–89	204	24	11.8	56	27.4	77	37.7	47	23.0
Total	1028	111	10.8	166	16.1	464	45.1	287	27.9
Women									
50–59	209	29	13.9	42	20.1	86	41.1	52	24.9
60–69	339	34	10.0	76	22.4	134	39.5	95	28.0
70–79	539	52	9.6	114	21.2	215	39.9	158	29.3
80–89	189	23	12.2	41	21.7	80	42.3	45	23.8
Total	1276	138	10.8	273	21.4	515	40.2	350	27.6

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[*] Shekelle RB, Dyer AR, Stamler J: Coffee, serum total cholesterol and risk of coronary heart disease in a population of employed, middle-aged men: The Western Electric Study. Presented at a Workshop on Coffee, Plasma Lipids and Coronary Heart Disease, Goteborg, Sweden, May 8–10, 1989.

TABLE 2—Age and Sex-Adjusted Means and Standard Errors of CHD Risk Factors and Proportions of Smokers, Exercisers, and Estrogen Users by Coffee Consumption, Rancho Bernardo 1984–87

	No Coffee mean (se) (n = 249)	Decaffeinated Only mean (se) (n = 439)	Caffeinated Only mean (se) (n = 979)	Both mean (se) (n = 637)
Age (sex-adjusted)	68.94 (.590)	71.79 (.444)	70.15 (.298)	70.68 (.369)
Obesity (kg/m²)	24.99 (.230)	25.39 (.175)	25.11 (.116)	25.09 (.144)
Alcohol (# drinks in past week)	4.08 (.495)	6.19 (.374)	7.08 (.250)	5.99 (.309)
Dietary Fat and Cholesterol index	33.97 (.388)	34.02 (.293)	35.86 (.195)	35.26 (.242)
Night Sleep (hours)	7.17 (0.79)	7.20 (.059)	7.27 (0.40)	7.23 (.050)
Smokers (% current)	5.40 (8.71)	11.16 (13.0)	15.05 (14.9)	11.69 (13.0)
Exercisers (% current)	82.49 (35.1)	82.07 (34.6)	80.23 (34.4)	80.91 (34.5)
Estrogen (women only)	26.34 (27.2)	34.11 (31.0)	27.17 (28.2)	27.54 (27.8)

TABLE 3—Age and Sex-Adjusted Means and Standard Errors of CHD by Consumption of Caffeinated Coffee, Rancho Bernardo, 1984–87

	Cups per Day			
Mean (SE)	1–2 (n = 216)	3–4 (n = 447)	≥5 (n = 953)	p-value (trend)
Age (sex-adjusted)	71.49 (.293)	69.62 (.428)	67.01 (.617)	<.01
Obesity	24.89 (.116)	25.53 (.169)	25.45 (.246)	.04
Alcohol (past week)	6.44 (.261)	7.08 (.380)	7.14 (.552)	.26
Dietary Fat and Cholesterol index	35.29 (.198)	36.46 (.288)	35.61 (.419)	.49
Night Sleep (hours)	7.24 (.041)	7.26 (.059)	7.32 (.086)	.42
Smokers (% current)	10.19 (12.3)	15.31 (15.11)	23.73 (19.2)	<0.1
Exercisers (% current)	82.38 (34.9)	80.27 (34.5)	72.45 (33.1)	<.01
Estrogen (women only % current)	25.80 (27.0)	30.63 (30.0)	25.63 (28.3)	.06

beverages, particularly coffee, showed a negative correlation with treadmill exercise capacity which persisted after cigarette smoking and other variables were partitioned out.

Coffee consumption could be an important link in the association between hours of nightly sleep and mortality. Karacan, *et al.*,[22] reported sleep disturbances in coffee drinkers while several other investigations of health-related practices have addressed the association between reported hours of nightly sleep and subsequent all cause mortality.[23–26] In this cohort, hours of nightly sleep were no different in coffee drinkers compared to non-drinkers and sleep hours did not vary with amount of caffeinated coffee consumed in either men or women. It is possible that older adults who found that caffeinated coffee disturbed their sleep patterns had discontinued its use.

In this cohort, 44 percent of men and 49 percent of women drank at least one cup of decaffeinated coffee per day. Behavior patterns in those who drank only decaffeinated coffee were more similar like those of caffeinated coffee drinkers than non-coffee drinkers. These findings suggest that decaffeinated coffee consumption should be considered in studies of the relationship beween coffee drinking and disease whenever a significant proportion of the study population drinks decaffeinated coffee.

In this study, coffee drinking was determined at the same time as all other behaviors, although the period of use for different behaviors varied from current to usual. Others have shown that current diet is a strong predictor of past or usual diet,[27] and we have found a strong correlation between usual and recent alcohol intake in this cohort. It seems unlikely that differences in the period of use explain the observed associations, unless one postulates that drinkers of coffee (with or without caffeine) have a better memory for unhealthy behavior than non-drinkers.

The diet data reported here were not directly validated; we know of no way to confirm dietary fat and saturated/unsaturated fat ratios short of direct observation or a fat biopsy.[28] Indirect validation of both food frequency and 24-hour diet recall was obtained in this cohort by demonstrating expected diet-disease associations. We have reported the expected cross-sectional associations of sodium, calcium, and potassium with blood pressure,[29,30] and inverse associations of fiber with heart attack;[31] potassium with stroke[32] and calcium with hip fracture in prospective studies.[33] Other attributes were also indirectly validated: smoking was inversely associated with HDL and directly with waist-hip-ratio;[34] alcohol was associated directly with waist-hip-ratio, aspartate aminotransferase and HDL;[35] exercise was positively associated with HDL and inversely associated with obesity, pulse rate, and plasma insulin.[36] Any misclassification of self-reported behaviors would be expected to reduce these associations, and not have caused them.

Because caffeinated and decaffeinated coffee consumption were both associated with more atherogenic behaviors, studies which seek to demonstrate the relationship between coffee consumption and morbidity or mortality must consider these other potential behavioral risk factors in the coffee-disease association. The concordance of heavy coffee intake with several lifestyles known to be related to heart disease risk implies a high level of potential confoundng in studies linking coffee consumption to hyperlipidemia and/or coronary heart disease.

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Validation of Mothers' Reports of Dietary Intake by Four to Seven Year-Old Children

CHARLES E. BASCH, PHD, STEVEN SHEA, MD, REBECCA ARLISS, PHD, ISOBEL R. CONTENTO, PHD, JILL RIPS, MPHIL,
BERNARD GUTIN, PHD, MATILDE IRIGOYEN, MD, AND PATRICIA ZYBERT, PHD

Abstract: The validity of mothers' recall of four to seven year-old children's diet was assessed among 46 first generation Latino immigrant families from the Dominican Republic by comparing intake recalled by the mother to unobtrusive home observations of children. Correlations were moderate to high for calories and for most nutrients. There were no differences in mean intake of total calories or in intake of most macronutrients and micronutrients assessed. At least two-thirds of the children in the lowest (or highest) quintile based on home observation were correctly classified into the lowest or second lowest (or highest) quintiles based on mother's reports for calories and most nutrients. For all food items that were both observed and reported, 51 percent of reported portion sizes were equivalent to observed portion sizes, 15.5 percent were smaller, and 33.5 percent were larger. There was fair to good agreement on the number of food items eaten, with the exception of vegetables. Mothers' recall appears to be useful for classifying children by intake of calories, macronutrients and micronutrients, but provides a somewhat less accurate measure of actual foods eaten, portion sizes, and nutrient levels consumed. (*Am J Public Health* 1990; 80:1314–1317.)

Introduction

Valid measurement of diet is essential to improve understanding of the role of diet in the etiology of diseases and for evaluating dietary interventions. The 24-hour dietary recall is probably the most widely used instrument in surveys and studies of food intake in the United States,[1–7] and its validity has been investigated in adults by comparison with weighed intakes,[8,9] direct observation,[10–12] and other dietary measurement methods.[13–15] Two recently published studies[16,17] have reported on the validity of recall by adults who cook and serve children's meals, usually mothers. No studies of this kind have been conducted in Latino families with young children or have used unobtrusive observation of children in their homes as a validation criterion. Nor have any previous studies assessed the validity of parental recall of children's nutrient intake using regression methods to adjust for caloric intake, permitting estimation of nutrient values that remove the effect of caloric intake.

We assessed the validity of mothers' recall of four to seven year-old children's dietary intake of total calories and 17 nutrients and of portion sizes and food items aggregated within food groups among first generation Latino immigrant families from the Dominican Republic by comparing intake recalled by the mother to unobtrusive home observations of children.

Methods

Subjects and Setting

Subjects were drawn from a sample of four to seven year-old children and their mothers participating in the Columbia University Study of Childhood Activity and Nutrition, a longitudinal study of young children. The total sample comprises families residing in a primarily Latino, densely populated, low-income neighborhood in northern Manhattan, New York City. Eligibility required that children be generally healthy with no special dietary prescriptions or restrictions.

A total of 46 Dominican families (mean age of children = 5.4 years, S.D. = 0.8) participated in the validation study. There were 18 males (mean age = 5.0 years, S.D. = 0.78 years) and 28 females (mean age = 5.7 years, S.D. = 0.78 years). All parents in the validation study were mothers (mean age = 33.8 years, S.D. = 4.6). Thirty-one of the children lived in single parent families.

Measurements

Home observation: Home observations were conducted between May 1987 and May 1989 by trained observers, fluent in Spanish, on the day before a regularly scheduled study visit in which the mother was to report her child's dietary intake for the preceding 24-hour period. The time period of the home observation ensured observation of the main meal (usually 4:00pm to 8:00pm). Parents were told that the observer was interested in the child's activities but were not informed about the specific aims of the study.

The observer followed the child and recorded the foods the child ate, the amounts, and the time. When necessary the observer asked the child's mother what the child was eating. Prior to data collection, observers were trained and tested in visually estimating portion sizes. Data were recorded on an instrument adapted from the Health and Nutrition Examination Survey.[7]

Twenty-four hour dietary recall: Different individuals conducted the interviews and home observations. On the day following the home observation the mother was asked to describe the food items eaten by the child during each eating occasion, the time of day, and portion sizes. Three dimensional food models, food packages, glasses, cups, spoons, and other utensils, and questions about food preparation methods and brand names of foods were used to increase accuracy. The sections of the recall corresponding to the time when the observer was present in the home were excerpted for analysis. For both observed dietary intake and 24-hour recall data, foods and portion sizes were converted to nutrient intakes using a dietary analysis program,[18] which used published nutrient data from the United States Depart-

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From the Center for Health Promotion, Teachers College, Columbia University (Basch, Contento, Rips, Gutin, Zybert); Department of Medicine and School of Public Health, Epidemiology, Columbia University College of Physicians & Surgeons (Shea); Health Promotion Program, New York City Police Department (Arliss); and Department of Pediatrics, Columbia University College of Physicians & Surgeons (Irigoyen). Address reprint requests to Charles E. Basch, Ph.D. Center for Health Promotion, Box 114, Teachers College, Columbia University, 525 West 120th Street, New York, NY 10027. This paper, submitted to the Journal December 1, 1989, was revised and accepted for publication April 26, 1990.

ment of Agriculture.[19,20] Supplemental nutrient values were obtained from other sources.[21]

Statistical Analysis

Nutrient intake values were adjusted for average caloric intake using regression techniques.[22] The natural logarithm transformation was used to improve normality for intake of nutrients but untransformed values are reported in the tables. Nutrient analyses were also performed using unadjusted values. Results were the same and only adjusted values are reported. Pearson product-moment correlation coefficients were used to estimate consistency between observed and reported calorie and nutrient intake. Paired t-tests were used to evaluate differences between means. We also examined the proportion of children correctly classified into quintiles of nutrient intake.[22] Consistency between observed and reported portion size was calculated within nine food groups comprising the 200 food items that were both observed and reported. Portion sizes were defined as equivalent if they were within one and one-half times the smaller amounts. The proportions of mothers who reported fewer, the same, or more items than observed within food groups were calculated, and the Z test[23] was used to assess differences in proportions.

Results

Correlations were moderate to high for most nutrients, although mean nutrient intakes derived from mothers' reports tended to be higher than those derived from observations (Table 1). Paired t-tests indicated no significant differences for total calories or macronutrients, except protein. For micronutrients, there were differences for cholesterol, calcium, iron, thiamine, riboflavin, and niacin. At least two-thirds of the children in the lowest quintile based on home observation were correctly classified into the lowest or second lowest quintiles based on mother's reports for all 17 nutrients except carbohydrates, calcium, and riboflavin (Table 2). Fewer children were correctly classified into the highest two quintiles for fats and cholesterol. For micronutrients, mother's reports were about as good as or better at high levels of intake than they were at low levels of intake. For all nutrients, mean observed intake in the highest quintile was at least three times as high as that in the lowest quintile, so that correct classification has practical significance.

Overall, for all 200 food items that were both observed and reported, 51 percent of reported portion sizes were equivalent to observed portion sizes, 15.5 percent were smaller, and 33.5 percent were larger (Table 3). When mothers made errors in reporting portion sizes they tended to overestimate for all food groups except butter/fats/oils. For all foods combined the proportion underestimated was significantly lower than the proportion over-estimated. The small numbers of dairy products (n = 4) and eggs consumed (n = 1) precluded meaningful analysis for these food groups.

There was fair to good agreement on the number of food items eaten, with the exception of vegetables (Table 4). In all nine food groups mothers tended to report fewer items eaten than observed. The practical significance of the percent agreement depends in part on the overall frequency of consumption. The proportion of mothers who agreed with the observer that items were not eaten is thus included in Table 4. For nine out of 10 of the most frequently eaten foods, observed and reported frequencies were identical (data not shown).

Discussion

Ability to measure children's absolute nutrient intake (e.g., 300 milligrams of cholesterol per day) as well as macronutrient intake as a proportion of absolute calories is important since some nutritional guidelines are stated in these terms.[24–26] Adjustment for caloric intake is necessary when assessing the association between particular nutrients and disease risk or the causal role of nutrients in disease processes.[27–29] The regression method of adjustment recommended by Willett and used in this study[22] permits the estimation of nutrient values while removing the effect of total caloric intake. This is not achieved by dividing nutrient values by total caloric intake.[30]

TABLE 1—Means, standard deviations (SD), Pearson correlation coefficients, and 95 percent confidence intervals for observed and reported nutrients and calories (data based on reports from 46 Dominican mothers residing in Northern Manhattan and corresponding home observations of children's diet, 1987–1989)*

	Observed mean		Reported mean			
	Mean	(SD)	Mean	(SD)	r	95% CI
Fat (g)	12.2	(5.4)	12.9	(7.1)	.52	(.27–.70)
Saturated fat (g)	3.0	(2.5)	3.3	(2.1)	.38	(.10–.60)
Monounsaturated fat (g)	4.1	(2.6)	4.0	(2.3)	.41	(.14–.63)
Polyunsaturated fat (g)	3.9	(2.6)	4.1	(3.0)	.58	(.35–.75)
Cholesterol (mg)	53.6	(84.6)	83.2	(230.5)	.65	(.44–.79)
Protein (g)	16.1	(7.1)	19.7	(8.0)	.50	(.25–.69)
Carbohydrates (g)	59.9	(14.3)	59.8	(16.0)	.48	(.22–.68)
Sodium (mg)	335.9	(302.0)	396.6	(698.4)	.27	(−.02–.52)
Potassium (mg)	566.5	(280.4)	656.9	(546.7)	.54	(.30–.72)
Calcium (mg)	63.0	(57.2)	90.6	(110.7)	.28	(−.02–.52)
Phosphorus (mg)	185.5	(64.9)	230.8	(108.3)	−.10	(−.38–.20)
Iron (mg)	3.0	(1.1)	3.6	(1.3)	.82	(.70–.90)
Vitamin A (IU)	1121.6	(1561.2)	2354.4	(5933.5)	.66	(.45–.80)
Thiamine (mg)	0.28	(0.11)	0.32	(0.12)	.68	(.48–.81)
Riboflavin (mg)	0.22	(0.10)	0.36	(0.43)	.52	(.26–.70)
Niacin (mg)	4.1	(2.2)	4.7	(2.3)	.72	(.55–.84)
Vitamin C (mg)	33.0	(37.3)	29.2	(29.3)	.50	(.25–.69)
Calories	465.0	(229.5)	507.0	(258.9)	.71	(.53–.83)

*Nutrient intakes were adjusted using residuals from regression models with caloric intake as the independent variable and nutrient intakes as the dependent variables (22).

TABLE 2—Comparisons of observed and reported nutrient and caloric intake based on joint classification by quintiles (data based on reports from 46 Dominican mothers residing in Northern Manhattan and corresponding home observations of children's diet, 1987–1989)*

	Joint classification by quintiles					
	Lowest quintile on observation			Highest quintile on observation		
	Lowest quintile on recall (%)	Lowest 2 quintiles on recall (%)	Highest quintile on recall (%)	Highest quintile on recall (%)	Highest 2 quintiles on recall (%)	Lowest quintile on recall (%)
Fat (g)	56	89	0	78	89	11
Saturated fat (g)	33	67	0	63	63	13
Monounsaturated fat (g)	67	100	0	50	63	0
Polyunsaturated fat (g)	44	100	0	44	67	11
Cholesterol (mg)	67	100	0	11	44	22
Protein (g)	44	67	11	67	89	0
Carbohydrates (g)	44	56	11	56	100	0
Sodium (mg)	33	67	0	56	89	11
Potassium (mg)	44	78	0	44	67	0
Calcium (mg)	44	56	22	67	78	11
Phosphorus (mg)	56	67	11	56	78	0
Iron (mg)	80	90	0	56	89	0
Vitamin A (IU)	44	89	0	56	100	0
Thiamine (mg)	56	89	0	44	78	0
Riboflavin (mg)	44	56	11	56	78	0
Niacin (mg)	67	100	0	44	89	0
Vitamin C (mg)	56	78	0	56	67	11
Calories	67	100	0	44	89	0

*Nutrient intakes were adjusted using residuals from regression models with caloric intake as the independent variable and nutrient intakes as the dependent variables (22).

TABLE 3—Portion size reported versus observed for the 200 food items both observed and reported by food group (data based on reports from 46 Dominican mothers residing in Northern Manhattan and corresponding home observations of children's diet, 1987–1989)*

Food Group	Number of Foods	% Observed > Reported	% Observed = Reported	% Observed < Reported
Eggs	1	0	0	100.0
Dairy products	4	0	75.0	25.0
Butter/fat/oil	13	46.2	30.8	23.1
Sugar	17	11.8	58.8	29.5
Meat/fish/poultry	23	13.0	52.2	34.8
Fruits	22	9.1	59.1	31.8
Beans/nuts	36	16.7	38.9	44.4
Bread/grains/rice/cereal	40	15.0	52.5	32.5
Vegetables	44	13.6	56.8	29.5
Total**	200	15.5	51.0	33.5

*Portion sizes were defined as equivalent if they were within one and one-half times the smaller amount.

**Difference in proportions overreported − underreported = .18 (95% confidence intervals .08–.27)

TABLE 4—Percent of respondents reporting fewer, the same or more food items (vs observed) by food group (data based on reports from 46 Dominican mothers residing in Northern Manhattan and corresponding home observations of children's diet, 1987–1989)

	Items Reported (vs Observed)			
Food Group	% Fewer	% Same	% More	% Agree not Eaten
Eggs	4.3	95.7	0	93.5
Dairy products	8.7	84.8	6.5	76.1
Butter/fat/oil	21.7	71.7	6.5	39.1
Sugar	17.4	69.6	13.0	37.0
Meat/fish/poultry	15.2	73.9	10.9	34.8
Fruits	19.6	69.6	10.9	45.7
Beans/nuts	15.2	73.9	10.9	28.3
Bread/grains/rice/cereal	10.9	80.4	8.7	15.2
Vegetables	30.4	41.3	28.3	17.4

Our findings support the utility of mother's recall of children's diet for research needing accurate classification of subjects' nutrient intake according to an ordinal or interval level of measurement. The differences in mean levels of intake for protein and several of the micronutrient values raise concerns about the use of parental recall for testing differences in mean nutrient intake between groups. Our results also indicate substantial errors in estimating portion sizes of foods, particularly over-estimation, as well as overreporting and underreporting of specific food items, and thus raise questions about the use of mother's recall as a method of measuring these aspects of children's eating behavior. Studies that require precise and accurate estimates of mean nutrient intake within groups or descriptions of food items and portion sizes consumed may need to rely on direct observation.

The dietary variable of interest in chronic disease epidemiology is typical diet. Eating patterns vary by time of week and year,[31–33] so that a single 24-hour dietary recall will not represent typical diet. Mothers' recall of children's evening meal may differ from that of other meals, and the presence of the observer may influence mothers' menu for the day or her likelihood of remembering what she served. The main evening meal is important, however, as evidenced by the observed consumption of calories, fat and saturated fat, as well as other nutrients.

Results from this study are consistent in some respects with other reports of the utility of surrogate-reported dietary information. The studies in adults, which reached guarded conclusions about the utility of this strategy, assumed that subjects' reports were accurate and these reports were the criterion against which the surrogates' reports were tested.[34–36] A trained observer provided the criterion measure in our study. Klesges, *et al*,[16] performed home observations of 30 children ranging in age from 24 to 48 months in which the observer weighed and recorded all food eaten by the child during the day. Dietary recalls were taken from the parent later that same day. At the nutrient level findings were similar to those in our study. At the food level, Klesges, *et al*, found that parents made errors in estimating portion sizes but correctly recalled 96 percent of the food items the child ate, a level of accuracy higher than in our study. Collecting recall data from the parents on the same day as the observation, rather than on the day following the home observation (which is how the recall method is generally used), or having the observer weigh food prior to consumption by the child may have enhanced the accuracy of reporting.

Eck, *et al*,[17] conducted unobtrusive observations of 34 children four to nine years of age eating a single meal in a cafeteria with both of their parents. Dietary recalls were taken from family members individually and as a group on the following day. There were no significant differences between observed and reported mean intake of total calories, calories from total fat, protein, carbohydrates, or sugar, or in intake of cholesterol, sodium, iron or calcium. The mean correlation for these nutrients was 0.77 for fathers, 0.63 for mothers, and 0.86 for the mother, father and child as a group. Eck, *et al*, also examined overreporting and underreporting of nutrients by food groups and reported that while general agreement was high, over- and underreporting varied greatly. The correlations between reported and observed nutrient intakes were, in general, higher than in our study. This study was based on a single meal in a cafeteria setting eaten by middle to high socioeconomic status families,[17] while our study was done in the home of Latino subjects of lower educational attainment.

Our study, together with the investigations by Klesges, *et al*, and Eck, *et al*, indicate that mothers' reports of children's diet appear to be useful to measure children's intake of calories, macronutrients, and micronutrients. The recall gives a somewhat less accurate measure of actual foods eaten, portion sizes, and nutrient levels consumed. Additional research is needed to verify our results in other Latino subgroups and non-Latino populations, to improve understanding about factors that improve and impede mother's reports about children's diet, and to test strategies to improve the validity of these measures.

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Development and Validation of a Food Use Checklist for Evaluation of Community Nutrition Interventions

Alan R. Kristal, DrPH, Barbara F. Abrams, DrPH, RD, Mark D. Thornquist, PhD, Loreli DiSogra, EdD, Robert T. Croyle, PhD, Ann L. Shattuck, MS, RD, and Holly J. Henry, MS, RD

Abstract: We developed a new dietary assessment instrument, the Food Behavior Checklist (FBC), which measures food use related to adopting lower-fat and higher-fiber diets. The FBC is a simplification of the 24-hour diet recall that consists of 19 simple yes/no questions about foods consumed during the previous day. To develop the FBC, an expert committee generated a list of foods based on our intervention program, and we used focus groups and random-digit dialing pretests to refine the format and clarify items. To validate the FBC, we compared responses of 96 women on the FBC to information collected during a professionally administered 24-hour diet recall. For most items, agreement between the FBC and 24-hour recall, based on the kappa statistic, was good to excellent. Agreement was poor for items requiring detailed knowledge about food composition (e.g. high-fiber cereal). There was a trend to over-report general food categories (e.g. luncheon meats) but not specific food items (e.g. ice cream). In an embedded randomized study, we found that a set of introductory items designed to serve as a memory retrieval cue did not improve agreement between the FBC and 24-hour recall. These data provide preliminary evidence supporting the use of short checklist questionnaires on the previous day's food use as a means to assess diet at the group or community level. This relatively inexpensive and rapid measure can be used to inform the design of public health nutrition programs and as an evaluation tool in intervention research. (*Am J Public Health* 1990; 80:1318–1322.)

Introduction

There is an emerging consensus that dietary change interventions are a necessary component of public health programs to prevent cardiovascular diseases and cancer.[1] Although the details of recommendations differ, the principal dietary goals shared by leading American public health organizations[1–3] are to: 1) reduce fat intake from approximately 37 to 30 percent or less of total calories; and 2) increase the intake of nutrient- and fiber-rich fruits and vegetables from approximately two to five or more servings per day. Much recent research has focused on community-level interventions, such as mass media campaigns and supermarket programs, because public health professionals believe that these interventions will be the most cost-effecitve means to change dietary behavior.[4]

Both the design and evaluation of public health nutrition intervention programs have problems in assessing dietary behavior in target communities. Standard diet assessment instruments used for clinical and etiological research such as multiple-day diet records and food frequency questionnaires are lengthy, expensive, and difficult to administer to representative population samples. Even the 24-hour recall, the traditional survey interview method used to assess group-level dietary intake,[5] is problematic in community intervention research; it takes approximately 30 minutes to complete, it is sufficiently complex to require both professionally trained interviewers and rigorous standardization and analysis procedures, and it is expensive to administer, code, and analyze. Clearly, new approaches to dietary assessment are needed.

Research on dietary assessment methods appropriate for community-level nutrition intervention research has focused on three approaches. The first approach is based on ecological-level measures: supermarket warehouse orders,[6] supermarket sales,[7] supermarket characteristics (such as percent of meat cases offering chicken, turkey, and fish[8]) and cafeteria receipts.[9] A second approach is based on short dietary questionnaires designed to assess only a limited number of nutrients.[10,11] A third approach is based on questionnaires that assess only the dietary behavior targeted by the intervention program. Several examples of these instruments have been described, for example the MRFIT fat counter,[12] but little is known about the statistical characteristics of these measures when used as program evaluation tools.

In this report we describe a new instrument to assess group-level dietary behavior that is based on the approach of assessing specific target behavior, the Food Behavior Checklist (FBC). The FBC is a simplification of the 24-hour recall, consisting of simple yes/no questions about foods consumed in the previous day. It focuses on a recent and defined time period and may thus reflect more accurately true food use compared to recall of perceived "usual" diet.[13] The advantage of the FBC over the traditional 24-hour recall is that the FBC can be administered quickly by non-professional personnel, needs no coding before analysis, can be completed over the telephone, and can be easily tailored to a specific intervention program. It was developed originally as a component of the evaluation of the Nutrition and Cancer Prevention Program of the California Department of Health Services. The FBC was subsequently incorporated into the evaluation of the Kaiser Family Foundation's Community Health Promotion Grants Program.* Here we describe the instrument, its development, and a study to evaluate its validity.

Methods

Development of the Questionnaire

The goal of the Nutrition and Cancer Prevention Program is to change the consumption of foods which are

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From the Cancer Prevention Research Unit, Fred Hutchinson Cancer Research Center, Seattle, WA; (Kristal, Thornquist, Croyle, Shattuck, Henry); the Departments of Health Services (Kristal), Biostatistics (Thornquist), and Psychology (Croyle), all at University of Washington, Seattle; School of Public Health, University of California at Berkeley (Abrams); and the California Public Health Foundation, Sacramento (DiSogra).

Address reprint requests to Alan R. Kristal, DrPH, Cancer Prevention Research Unit, Fred Hutchinson Cancer Research Center, 1124 Columbia Street MP 702, Seattle, WA 98104. This paper, submitted to the Journal February 20, 1900, was revised and accepted for publication August 13, 1990.

* Wagner EH, Kehrer B. The Henry J. Kaiser Family Foundation Community Health Promotion Program. Evaluation design considerations. Presented at the annual meeting of the International Epidemiological Association, Helsinki, Finland, August 1987.

believed to contribute to the development of cancer—to increase the use of foods containing protective nutrients (vitamin C, carotenoids, and dietary fiber) and decrease the use of foods high in fat. The program goals were operationalized as: increasing the consumption of fruits, vegetables, whole grain/high fiber products and reduced-fat milk, and decreasing the consumption of full-fat dairy products, fried foods, ground beef and other high-fat meats, pastries, cakes, ice cream, and butter or margarine on bread and vegetables. We used these intervention goals to generate a list of foods for inclusion in the FBC.

The draft FBC was reviewed by an expert panel, including epidemiologists, nutritionists, a statistician, and a cognitive psychologist. A series of pre-tests was completed over the telephone, using respondents from a broad range of socioeconomic and age groups. Based on the results of these steps, an introductory section asking the respondent to recall which meals and snacks were consumed the previous day was added, several items were eliminated, and the questionnaire was reworded and reordered.

We then used focus groups to examine the relevance of the instrument to different ethnic groups (Black, Hispanic and Asian) and to men as well as women. The questionnaire was again reworded, and several ethnically specific foods were added as descriptions to the original food behavior items. Finally, we completed a full-scale, random-digit dial pilot survey of 299 adult residents of two California communities. From this pilot we compiled a final interviewer manual and code book (available from the authors) and completed minor rewording. The 19 items included in the instrument are given in the Appendix.

Validation Study

Detailed descriptions of the sampling scheme, participant demographic characteristics, and dietary assessment methods used in this study have been published.[11,14] Briefly, participants were 96 women ages 45 through 59 years who were members of a large health maintenance organization in Western Washington State. Dietary assessments included two, four-day diet records (at study entry and six months later) and a food frequency questionnaire (two months after study entry). Participants received no information about nutrition or their diets until all data were collected.

The validity study was completed by telephone three months after study entry. The interview consisted of: the FBC; a four-minute set of questions on usual dietary behavior; and a detailed, standardized 24-hour recall administered by a research nutritionist (HH or AS). The FBC used in the validation study differed slightly from the final FBC because pretesting had not yet been completed. These differences are given in the Appendix. The 24-hour recall was administered by working the participant through their previous day, using prompts to obtain a complete list of foods and asking questions to get specific information (brand of cereal, type of bread, etc). Nutritionists defined a high-fiber cereal as one with at least 3g of dietary fiber per serving; whole grain breads included all breads with the exception of white breads and caramel-colored breads made without whole grain flour. We considered this 24-hour recall to be a good criterion measure of foods eaten in the previous day because 24-hour recalls are the most commonly used method to obtain short-term, retrospective data on dietary intake.[5]

We also considered whether the use of introductory questions on the context of the previous day's food consumption would improve the validity of responses to the FBC. Studies of human memory have shown that context can serve as an effective retrieval cue for memories of personal events.[15] Furthermore, studies of survey respondents indicate that many persons first think about the circumstances surrounding a health event when they are asked to specify its date or frequency.[16] This suggests that specific, incidental events such as eating a bowl of cereal may be better remembered if the individual first thinks of the more easily remembered context such as having breakfast. Study participants were thus randomized into two groups: receiving or not receiving five introductory questions before beginning the FBC. These introductory questions were similar to probes used during a 24-hour recall, and asked about whether the participant ate breakfast, lunch, and dinner, whether they ate in a restaurant and, if so, in what type of restaurant (fast food/other).

Statistical Methods

We used two approaches to measure agreement between the FBC and the criterion measure, the 24-hour recall.

- The first approach was to compare the proportion of participants who reported each food behavior from the FBC to the proportion reporting the same behavior during the 24-hour recall. This comparison considers whether FBC and 24-hour recall would yield different community-level estimates of dietary behavior practiced in the previous day. We calculated confidence intervals around these differences using the exact binomial distribution.[17]
- The second approach to assess validity was to examine the proportion of individuals with the same responses from the FBC and the 24-hour recall. This analysis considers whether the FBC and 24-hour recall differ when measuring the behavior of individuals. We used the kappa statistic to assess agreement in this application.[18] In presenting the results, we indicated items for which expected chance agreement was over 90 percent because the kappa statistic in this case will be based on very few discordant pairs and thus be highly unreliable.[19]

Finally, we examined the association of items on the FBC with usual percent of energy from fat and usual fiber intake. To characterize usual dietary intake in these analyses we used the average of values from the food frequency questionnaire and eight days of diet records. We did not expect that all items on the FBC would be strongly associated with usual dietary intake because the use of a food in the previous day is not a reliable measure of usual food use.[5] However, we considered associations of FBC items with usual dietary intake as evidence of the content validity of the instrument.

Results

Table 1 gives the proportion of participants reporting use of 18 foods in the previous day from both the FBC and the 24-hour recall. Overall, the proportion of yes responses tended to be higher for the FBC than the 24-hour recall (sign test $p < 0.01$), although the 95 percent confidence intervals for these differences excluded zero for only five items: high-fiber cereal, fruit, a vegetable at lunch, a vegetable at dinner, and whole-grain bread. The percent of concordant responses and the corresponding kappa statistic for each item are also given in Table 1. Observed agreement was under 85 percent for only three items: high-fiber cereal; whole-grain breads; and butter or margarine on bread. Only two kappas, for high-fiber cereal and bacon or sausage, were under 0.60,

TABLE 1—Agreement between Food Behavior Checklist and 24-Hour Recall

Food Behavior Checklist Item*	Responded Yes: Food Behavior Checklist (%)	Responded Yes: 24-Hour Recall (%)	95 Percent Confidence Interval for Difference	Agreement: (%)	Agreement: Kappa
Cereal	50.0	50.0	(−1.9, 1.9)	97.9	0.96
High-fiber cereal	89.1	65.2	(12.0, 23.9)	76.1	0.37
Pastry	13.5	9.4	(0.0, 4.2)	95.8	0.80
Bacon or sausage	3.1	4.2	(−3.0, 2.3)	96.9	0.56**
Milk	65.3	65.3	(−0.2, 0.2)	91.6	0.81
Fruit	90.5	84.2	(1.2, 6.3)	93.7	0.72
Salad	52.1	45.8	(−6.8, 10.2)	85.4	0.71
Vegetable at lunch	34.4	24.7	(0.0, 13.1)	86.0	0.67
Vegetable at dinner	68.4	53.7	(8.7, 14.7)	85.3	0.70
Butter on vegetable	19.0	12.6	(−3.9, 9.6)	89.5	0.61
Dark bread	56.3	45.8	(2.4, 14.9)	83.3	0.67
Muffins	17.7	15.6	(−3.5, 6.5)	91.7	0.70
Butter on bread	37.5	33.3	(−4.9, 10.7)	83.3	0.64
Fried food	7.3	7.3	(−1.9, 1.9)	97.9	0.85
Cold cuts	6.3	6.3	(−2.3, 2.3)	95.8	0.64
Ground beef	13.5	13.5	(−2.3, 2.3)	95.8	0.82
Cakes	45.8	45.8	(−2.3, 2.3)	95.8	0.92
Ice Cream	14.6	14.6	(−1.9, 1.9)	97.9	0.92

*Full text of items in Appendix.
**Expected chance agreement >90%.
N = 96

a level that may be considered a cutpoint for fair agreement beyond chance. For the remaining 16 items, agreement was very good to excellent for 11 (kappa > 0.80) and good for five (kappa > 0.60).

In further analysis stratified by whether or not participants received the introductory questions to the FBC, there was no evidence that introductory questions improved the quality of responses (data available from the authors). The tendency for a higher proportion of participants to report food use in the FBC than in the 24-hour recall remained evident in both strata. This bias was not smaller among participants who received the introductory questions, nor was there a trend for kappas to be higher.

The associations of usual dietary intakes of fat and fiber (assessed from a food frequency and eight days of dietary records) with responses to the FBC are given in Table 2. With only a single exception (vegetable at dinner) the direction of association was consistent with the fat or fiber content of each food. Confidence intervals for differences in percent of energy from fat excluded zero for consumption of pastry, salad, butter or margarine on breads, and cold cut luncheon meats. Confidence intervals for fiber intake excluded zero for only cereal and fruit. These results did not differ when food use in the previous day was taken from the 24-hour recall.

Discussion

Overall, reports of previous day's food behavior related to fat and fiber intake appeared similar when assessed by either the Food Behavior Checklist or a 24-hour recall. There

TABLE 2—Mean Usual Percent of Energy from Fat and Fiber Intake, by Response to Food Behavior Checklist

Food Behavior Checklist Item*	Percent of Energy from Fat Response to FBC: "Yes"	Percent of Energy from Fat Response to FBC: "No"	95 Percent Confidence Interval for Difference	Fiber (g) Response to FBC: "Yes"	Fiber (g) Response to FBC: "No"	95 Percent Confidence Interval for Difference
Cereal	31.8	33.2	(−3.9, 1.0)	18.2	16.1	(0.0, 4.1)
High-fiber	31.9	31.5	(−5.7, 6.4)	18.6	15.3	(−1.4, 8.0)
Pastry	35.8	32.0	(0.2, 7.3)	16.7	17.2	(−3.6, 2.6)
Bacon or sausage	37.5	32.3	(−1.8, 12.2)	15.5	17.2	(−7.8, 4.4)
Milk	32.4	32.6	(−2.8, 2.4)	17.5	16.4	(−1.1, 3.4)
Fruit	32.3	34.9	(−6.8, 1.6)	17.6	13.2	(0.8, 7.9)
Salad	30.3	34.9	(−6.9, −2.3)	17.7	16.5	(−0.9, 3.3)
Vegetable at lunch	31.6	32.8	(−3.8, 1.5)	18.2	16.4	(−0.4, 4.1)
Vegetable at dinner	32.0	33.6	(−4.2, 1.1)	16.8	17.7	(−3.1, 1.4)
Butter on vegetable	34.6	32.0	(−0.5, 5.7)	17.4	17.1	(−2.5, 3.0)
Dark bread	31.8	33.4	(−4.1, 0.8)	18.0	16.1	(−0.2, 4.0)
Muffins	30.9	32.8	(−5.1, 1.3)	18.0	16.9	(−1.7, 3.8)
Butter on bread	34.4	31.4	(0.6, 5.5)	17.1	17.2	(−2.6, 2.4)
Fried food	34.2	32.4	(−2.9, 6.6)	16.5	17.2	(−4.8, 3.3)
Cold cuts	39.4	32.0	(2.5, 12.2)	14.3	17.3	(−7.4, 1.3)
Ground beef	34.2	32.2	(−1.6, 5.6)	15.9	17.3	(−4.5, 1.7)
Cakes	33.0	32.1	(−1.6, 3.4)	16.1	18.0	(−3.9, 0.3)
Ice cream	35.0	32.1	(−0.5, 6.4)	16.9	17.2	(−3.2, 2.7)

*Full text of items in Appendix.

was a general trend for overreporting food use on the FBC, but this did not appear to be a bias to report more socially desirable behavior. Both desirable (e.g. "vegetable at dinner" and "high-fiber cereal") and undesirable (e.g. "butter or margarine on bread" or "breakfast or snack pastry") behavior were reported more frequently on the FBC than the 24-hour recall.

We suspect that forward telescoping,[20] a term used in cognitive psychology to describe recall of a past event as more recent in time, may explain higher responses from the FBC. In this study, we observed telescoping for common or non-specific dietary behavior, such as vegetable and fruit eating; responses to more specific dietary behavior, such as use of cold cuts or ice cream, were identical from the FBC and 24-hour recall. This suggests that a checklist approach to dietary assessment should emphasize specific rather than general dietary behavior. For example, it may be better to read a list of specific vegetables or to ask "What vegetables did you eat with lunch" rather than to ask simply "Did you eat a vegetable with lunch?"

From a cognitive perspective, the FBC has a number of advantages over more complex instruments to assess dietary behavior. First, the question format is straightforward. This feature reduces variation in interviewer behavior and simplifies the response task of the participant.[21] Second, participants do not appear to find remembering the previous day's food behavior difficult. Finally, the use of a 24-hour period provides the respondent with a simple landmark. For questions on specific dietary behavior this may have reduced the telescoping errors typical of more long-term recall periods.[22]

There was also evidence that questions requiring high levels of nutritional knowledge did not produce valid data. For the questions on high-fiber cereal and whole-grain bread we determined that the source of disagreement between the FBC and 24-hour recall was due to the nutritionists' correct classification of the food based on brand name information obtained during the 24-hour recall. For example, participants may report that they are eating a high-fiber cereal when, based on published nutrient composition data, a serving of the cereal actually contained less than 3 grams of fiber per serving. In the final FBC, the "high-fiber" cereal question was modified to a list of specific high fiber cereals, eliminating the need for the respondent to know the fiber content of cereals they ate.

We had hypothesized that the introductory questions would serve as a memory retrieval cue and improve the accuracy and speed of responses. Contrary to our expectations, interviews with the five simple, introductory questions took an average of 1.4 additional minutes to complete ($p < 0.01$) and resulted in no meaningful improvements in agreement between the FBC and 24-hour recall. These results suggest that the FBC items did not pose a significant challenge to respondents' memory abilities.

Given the characteristics of participants in this study there are limits to generalizing these results to more representative population samples. Participants in this study were all women, interested in diet and health, well educated, and willing to comply with intensive measurement protocols. Evaluation of the FBC in other populations is now under investigation. The FBC has been used successfully during baseline assessments of 15 communities in the Kaiser Family Foundation's Community Health Promotion Grants Program, and we have found strong community-level associations between the FBC and measures of supermarket characteristics.[23] The FBC has also been incorporated into an ongoing random-digit dial health survey of Washington State residents, and validation studies from this survey are underway.

Data collected from the FBC lend themselves to two types of analyses. When data are collected at the beginning of a planned intervention, they can be used to document problems and formulate program goals. When data from an FBC are collected pre- and post-program implementation in intervention and control communities, they can be used to evaluate changes in the prevalence of food use attributable to an intervention program. Thus, if intervention program goals can be clearly formulated as increases or decreases in the prevalence of use of specific foods, the FBC may be a superior instrument to standard dietary questionnaires which attempt to measure total intake of nutrients.

In conclusion, there is preliminary evidence that a checklist questionnaire is a reasonable approach to community-level dietary assessment. We have demonstrated that a careful process of selecting questionnaire items can produce an instrument with both face and content validity, and answers to a checklist questionnaire are similar to those obtained from a 24-hour recall. The FBC may be an inexpensive and technically simple dietary assessment tool when group-level data on dietary behavior can be used to evaluate intervention programs or monitor dietary change.

APPENDIX

Food Items Included in Food Behavior Checklist

Cereal (hot or cold)
    a. If yes, was it a high fiber cereal?*
        a. If yes, was it any of the following? Any bran cereal, Raisin Bran, Wheatena, Shredded Wheat, Oatbran, Oatmeal or Fruit n Fiber?**
Breakfast or snack pastry such as donuts, croissants, danish, coffee cake, or sweet rolls, pan dulce or custard tarts
Bacon, sausage, or chorizo
Milk, including milk on cereal, chocolate milk, or cafe con leche
    a. If yes, was it**
        regular or whole milk?
        low-fat or 2% milk?
        skim, 1% or non-fat milk?
One or more servings of fresh fruit
Green salad
Vegetable, other than salad, at lunch
Vegetable, other than salad, at dinner
Butter, margarine, or creamy sauce on cooked vegetables
Dark bread, including whole wheat, rye, or pumpernickel
Whole grain or bran muffins
Butter or margarine on bread
Deep-fried food like french fried potatoes, fried chicken or fish, tacos, egg rolls, or dim-sum
Hot dog, salami, bologna, or other luncheon meat
Hamburger, meat loaf, tacos or other dish with ground beef
Cake, pie, custard, tart or cookies
Ice cream

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* Included in validity study only

** Included in final instrument only

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Food and Nutrient Intake Differences between Smokers and Non-smokers in the US

Amy F. Subar, PhD, RD, Linda C. Harlan, PhD, and Margaret E. Mattson, PhD

Abstract: Data from the Second National Health and Nutrition Examination Survey were analyzed to determine food and nutrient intake differences between current smokers (also categorized as light, moderate, and heavy smokers) and non-smokers. Smokers in several age-race-sex categories have lower intakes of vitamin C, folate, fiber, and vitamin A than non-smokers, and intake tended to decrease as cigarette consumption increased, particularly for vitamin C, fiber, and folate. Smokers were less likely to have consumed vegetables, fruits (particularly fruits and vegetables high in vitamins C and A), high fiber grains, low fat milk, and vitamin and mineral supplements than non-smokers. A negative linear trend was found between smoking intensity and intake of several categories of fruits and vegetables. These data suggest that the high cancer risk associated with smoking is compounded by somewhat lower intake of nutrients and foods which are thought to be cancer protective. (*Am J Public Health* 1990; 80:1323–1329.)

Introduction

Cigarette smoking is the biggest single health habit that contributes to preventable chronic disease in the United States especially cancer of the lung, larynx, oral cavity, and esophagus as well as coronary artery disease and emphysema.[1] Other cancers, such as bladder, kidney, pancreas, cervix, and stomach are also associated with cigarette consumption. Cigarette use in the US has been estimated to be a causative factor in 30 percent of all cancer deaths and in 87 percent of all lung cancer deaths.[2] It has also been estimated that diet is a causative factor in 35 percent (with a possible range of 10–70 percent) of all cancer deaths. While smoking, food, and alcohol consumption each are lifestyle practices associated with cancer risk, when they occur together some authors assert they act synergistically to produce risks which are greater than the sum of their parts.[3]

Compared to non-smokers, smokers consume more alcohol and coffee, take more risks, have more sexual partners, exercise less, sleep less, take vitamin and mineral supplements less, and skip breakfast more.[4–11] Furthermore, several of these poor health habits are more likely to occur with increasing average number of cigarettes smoked per day.[8]

Among smokers, those with either higher dietary intakes of vitamin and/or carotene or higher serum carotenoids are at reduced risk for lung cancer,[12–25] and several studies have examined the vitamin A, carotenoid, or energy intake differences between smokers and non-smokers.[26–30] A recent study showed an inverse association between intake of vitamin C and smoking.[31] Only one study[32] of 45–59 year old men enrolled in a prospective heart disease trial in Wales, however, examined total nutrient intake differences between smokers and non-smokers, and smokers were found to have lower intakes of most nutrients.

Understanding the dietary patterns of those who smoke compared to those who do not smoke is important for health professionals seeking to assess disease risk and educate individuals regarding healthy lifestyles. We used data from a large nationally representative survey, the Second National Health and Nutrition Examination Survey (NHANES II),[33] conducted from 1976–80, to determine whether smokers significantly differ from non-smokers in both nutrient and food intake.

Methods

NHANES II used a probability sample of the civilian non-institutionalized US population ages 6 months to 74 years. The complex sample design and sample weights are described in detail elsewhere.[34] Mobile examination centers employing trained teams of physicians, nurses, dietitians, and technicians traveled to and collected data from 64 sites throughout the country.

NHANES II 24-hour dietary recall data from 11,260 adults ages 19–74 years were used to assess dietary intake. For these analyses, races other than Black or White were excluded because of small sample sizes. Also excluded were individuals (N = 280) with imputed, surrogate, or unreliable 24-hours recall data. Trained interviewers collected dietary intake data from the day before the interview. Three-dimensional food portion models were used to help individuals accurately report portion sizes.[35] Dietary data were coded by the interviewers within 72 hours of data collection and analyzed for nutrient content using current nutrient values. Nutrient values for folate and dietary fiber, however, were not available on NHANES II tapes and were added later by Subar, *et al*,[36] and Lanza, *et al*,[37] respectively.

A smoking history was obtained from all subjects. Individuals were placed into categories of current, former, and never smokers, and current smokers were divided into categories of light, moderate, and heavy smokers based upon reported average use of 1–14, 15–24, or ≥25 cigarettes/day, respectively. Initial analyses showed few differences in nutrient or food intakes from 24-hour recall data between never and former smokers and therefore they were combined, creating a group of current non-smokers. Individuals who were not cigarette smokers but smoked either pipes (N = 63) or cigars (N = 140) <6/week, and <3/week, respectively, were included as non-smokers because usage was considered too infrequent to warrant exclusion. Frequent pipe or cigar users (N = 398) were excluded from the analyses.

Nutrient intake differences between current smokers and non-smokers, and among current non-smokers and light, moderate, and heavy smokers in the 24-hour recall data set were examined by race, sex, and broad age groups (19–29, 30–54, and 55–74 years). Differences in food intake were determined using food groups developed by Patterson and Block.[38] The food groups created were based upon the

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From the National Cancer Institute, Division of Cancer Prevention and Control, National Institutes of Health. Address reprint requests to Amy F. Subar, PhD, RD, Division of Cancer Prevention and Control, NCI, NIH, 9000 Rockville Pike, Executive Plaza North 313, Bethesda, MD 20892.

interim dietary guidelines by the National Research Council's report, *Diet, Nutrition and Cancer*,[39] and those of the American Cancer Society.[40] An inclusive "all vegetables" category was developed as well as the following vegetable subcategories: cruciferous vegetables (i.e. broccoli, cauliflower, brussels sprouts, cabbage), deep yellow vegetables (carrots, winter squash, sweet potatoes), fiber-rich vegetables (dried peas, beans, corn), and "garden vegetables" (all vegetables except white potatoes, dried peas and beans, and salad). Fruits and vegetables were also subdivided into categories of those high in either vitamin A (i.e. carrots, spinach, cantaloupe) or vitamin C (i.e. oranges, grapefruit, broccoli). A complete listing of foods in each category has been published.[38] Regardless of portion size or numer of servings, individuals were described as having consumed or not consumed food(s) within each food group based upon 24-hour recall data. Preliminary analyses of food group intake indicated that such coding yielded few differences by sex and, therefore, individuals were categorized by age and race only.

In addition to univariate analyses described above, multivariate analyses of nutrient and food intake were performed adjusting for age, kcal/daily, poverty index ratio or PIR (an income status index which takes into account household income, family size, and the income necessary to maintain a family at the time the study was conducted), and sex. These analyses were performed separately for Whites and Blacks because of the known racial differences in consumption of foods and nutrients.[41]

Separate analyses by sex-race categories were performed to determine whether smoking category based upon average daily cigarette consumption, as measured in half packs of cigarettes consumed, was linearly related to intake of nutrients or food groups. Each nutrient or food group, acting as a dependent variable in the model which adjusted for age, examined the effect of half packs of cigarettes consumed. For purposes of these analyses, quartiles of nutrient intake by race-sex groups were used.

Statistical analyses were performed using SAS[42] with appropriate population weights permitting inferences about the total adult US Black and White population. Standard errors were calculated using SESUDAAN,[43] a program designed to account for the complex survey design. The SURREGR[44] program for complex samples was used in multivariate analysis for the White population, but because of the lack of dispersion of Blacks throughout strata and primary sampling units (PSU), an unweighted regression analysis was used for Blacks. Logistic regressions were performed using the RTIlogt program.[45] This application uses the Taylor expansion series to estimate the variances.

Results

Food Group Intake Differences

Analyses by age, race, sex, and current smoking status showed a larger proportion of Black men (55 percent) smoked during the period of the NHANES II study than in any other sex-race category, followed by White men (43 percent), White women (34 percent), and Black women (34 percent). In all age-race-sex groups, approximately 50 percent of all smokers are 30–54 years of age, while approximately one-third and 15 percent are ages 19–29 and 55–74 years, respectively.

Tables 1 and 2 present univariate analyses of food group intake of smokers and non-smokers by age for Whites and Blacks, respectively. With few exceptions, intake of vegetable categories except "starchy fibrous" is moderately higher among non-smokers as compared to smokers in all age-race groups. Multivariate analyses which adjusted for age, kcal/day, PIR, and sex confirmed these results among Whites. Both univariate and multivariate analyses show that non-smokers of both races have a strikingly higher intake of fruits compared to smokers with percentage differences as high as 19 percent (unadjusted) for 19–29 and 30–54 year old Whites, and 55–74 year old Blacks. For the categories of high fiber grains and fruits or vegetables high in either vitamin C or vitamin A, non-smokers in all age-race categories have a greater intake than smokers, a finding confirmed after multivarate adjustment. For high fiber grains, however, it should be noted that Blacks have much lower intakes than Whites with the highest intakes among Black non-smokers equal to or less than the lowest intakes of White smokers.

TABLE 1—Weighted Percent of White Subjects Consuming Food(s) Over a 24-Hour Period (NHANES II) by Age and Smoking Status

	Whites 19–29 years				Whites 30–54 years				Whites 55–74 years			
Food Group	Non-Smoker	95% CI[a]	Smoker	95% CI	Non-Smoker	95% CI	Smoker	95% CI	Non-Smoker	95% CI	Smoker	95% CI
All Vegetables	84.4	82.6,86.1	79.0	76.7,81.3	85.9	84.4,87.4	81.0	79.2,82.9	86.3	84.7,87.9	82.9	79.9,86.0
Garden	49.7	47.0,52.4	42.7	39.3,46.1	54.2	51.4,57.0	48.5	46.4,50.5	58.4	55.9,61.0	51.9	48.3,55.5
Deep Yellow	9.2	7.4,11.1	6.3	4.4,8.1	11.6	9.8,13.4	8.7	6.8,10.6	13.7	12.2,15.2	11.0	8.8,13.2
Cruciferous	15.3	13.4,17.2	12.0	9.9,14.1	20.0	18.5,21.5	15.9	13.4,18.3	21.2	19.3,23.0	18.4	16.3,20.4
Starchy Fibrous	18.8	16.0,21.5	18.8	15.9,21.7	19.1	17.2,21.1	18.2	16.2,20.2	18.0	15.4,20.5	17.9	15.2,20.6
Salad	48.6	45.3,52.0	39.4	35.7,43.0	50.5	48.4,52.6	46.7	43.6,49.9	49.7	47.3,52.1	46.1	42.4,49.8
All Fruits	59.3	56.4,62.3	39.6	35.7,43.5	63.9	61.7,66.1	45.4	42.4,48.5	76.8	74.8,78.7	63.3	60.0,66.5
Fruits/Vegetables												
High in Vitamin A	18.4	15.7,21.1	13.3	11.5,15.1	23.2	21.2,25.2	15.9	13.8,18.1	26.1	24.1,28.1	20.8	17.9,23.7
High in Vitamin C	25.6	22.5,28.7	19.5	16.7,22.3	31.6	29.2,34.1	20.4	17.9,22.9	36.5	33.2,39.7	27.6	24.1,31.1
High Fiber Grains	16.6	13.1,20.0	9.9	6.8,12.9	18.1	15.4,20.9	11.1	8.4,13.7	25.5	23.1,27.9	20.0	16.1,23.8
Red Meat	59.5	56.4,62.6	60.5	57.5,63.4	56.0	53.8,58.3	56.8	54.1,59.5	50.2	48.2,52.1	50.1	46.1,54.1
Poultry/Fish	31.8	29.0,34.6	29.3	26.7,31.9	33.0	30.5,35.5	28.7	25.7,31.8	34.8	32.6,37.1	30.4	26.8,33.9
Bacon/Luncheon Meat	36.2	33.6,38.9	41.2	37.6,44.7	42.4	39.6,45.2	46.6	43.1,50.0	41.4	38.6,44.2	44.9	40.9,49.0
Whole Milk	41.8	38.5,45.0	47.6	43.7,51.4	40.5	37.8,43.1	43.3	39.7,46.9	39.6	36.7,42.6	46.0	41.8,50.3
Low Fat Milk	30.2	26.2,34.3	22.0	18.1,25.8	28.7	25.9,31.6	20.3	17.0,23.7	34.5	31.7,37.4	24.2	20.2,28.3
Use Supplements[b]	24.3	22.0,26.6	15.6	13.3,17.9	24.9	22.4,27.3	20.8	18.0,23.6	34.6	32.1,37.1	29.3	25.9,32.8

a) CI = Confidence interval.
b) Data based upon reported "regular" usage of vitamin/mineral supplements.

TABLE 2—Weighted Percent of Black Subjects Consuming Food(s) Over a 24-Hour Period (NHANES II) by Age and Smoking Status

	Blacks 19–29 years				Blacks 30–54 years				Blacks 55–74 years			
Food Group	Non-Smoker	95% CI[a]	Smoker	95% CI	Non-Smoker	95% CI	Smoker	95% CI	Non-Smoker	95% CI	Smoker	95% CI
All Vegetables	77.5	69.4,85.6	79.6	72.2,87.0	77.8	72.2,83.4	70.6	63.1,78.1	80.2	74.2,86.1	70.6	62.5,78.6
Garden	43.8	34.3,53.3	42.1	33.3,50.9	52.5	41.8,63.3	41.2	33.1,49.3	58.6	50.1,67.1	43.6	34.9,52.4
Deep Yellow	6.3	2.6,9.9	6.4	2.4,10.4	8.7	2.3,15.0	7.4	4.0,10.8	17.8	12.8,22.8	6.5	3.0,10.1
Cruciferous	21.1	12.5,29.7	18.9	12.0,25.9	23.7	13.4,33.9	21.3	15.1,27.5	28.4	20.7,36.2	22.6	11.8,33.4
Starchy Fibrous	22.5	18.3,26.7	18.3	11.4,25.3	17.6	12.7,22.5	13.2	8.5,17.9	18.8	11.0,26.5	27.0	17.8,36.2
Salad	34.5	27.7,41.3	25.7	18.7,32.6	26.7	19.8,33.7	26.4	21.5,31.4	25.1	17.7,32.5	17.1	9.6,24.6
All Fruits	62.0	54.4,69.5	47.8	38.0,57.5	61.2	53.4,69.0	49.2	40.6,57.8	64.0	56.0,72.0	45.3	37.2,53.5
Fruits/Vegetables												
High in Vitamin A	21.6	13.6,29.7	19.3	12.4,26.2	26.2	16.1,36.2	25.0	17.6,32.4	38.6	30.9,46.3	22.7	12.2,33.2
High in Vitamin C	34.7	26.9,42.6	30.7	23.8,37.6	33.5	25.3,41.6	29.7	23.9,35.5	44.7	35.3,54.0	32.7	22.6,42.8
High Fiber Grains	8.9	5.0,12.8	4.9	1.6,8.2	11.6	6.5,16.6	6.4	3.0,9.8	10.9	6.9,14.8	5.8	0.0*,11.9
Red Meat	58.2	49.6,66.8	56.3	49.7,62.8	44.8	38.8,50.8	45.9	38.4,53.4	35.6	28.7,42.4	35.1	27.2,42.9
Poultry/Fish	47.4	36.5,58.4	41.2	31.7,50.7	55.3	48.5,62.1	48.2	40.3,56.2	52.2	44.4,60.1	47.8	36.0,59.5
Bacon/Luncheon Meat	40.4	34.5,46.2	47.8	37.1,58.6	41.8	35.6,48.0	48.2	41.3,55.2	45.7	39.6,51.9	52.8	43.8,61.8
Whole Milk	41.4	34.8,48.1	35.0	28.2,41.7	27.1	18.0,36.2	32.9	24.6,41.2	33.8	24.7,42.9	33.6	24.4,42.8
Low-Fat Milk	3.5	0.3,6.8	2.7	0.3,5.2	18.7	11.9,25.5	4.5	1.1,7.9	15.7	9.1,22.3	10.2	4.3,16.0
Use Supplements[b]	13.3	8.9,17.7	13.6	7.9,19.2	23.8	16.3,31.3	12.4	8.5,16.3	2.0	14.5,25.5	16.6	9.1,24.0

a) CI = Confidence interval.
b) Data based upon reported "regular" usage of vitamin/mineral supplements.
*Lower limit truncated to 0.0 from a negative value.

Univariate analyses (Tables 1 and 2) showed that non-smokers compared to smokers tended to consume equal amounts of red meat, but more poultry/fish and skim milk, and less bacon/luncheon meats, and whole milk. These findings were confirmed in the multivariate model. In addition, among Whites and most Blacks, non-smokers were more likely to take vitamin/mineral supplements. Blacks, however, were low supplement users overall.

Nutrient Intake Differences

Tables 3 and 4 present, for Whites and Blacks, respectively, univariate analyses of mean intake data for energy, seven nutrients, dietary fiber, and cholesterol by age, sex, and current smoking status. Analyses were also performed for calcium, iron, thiamin, riboflavin, niacin, oleic acid, and linoleic acid, but results showed few or no differences by smoking status (data available upon request to author). As found in past research of the NHANES II data set,[27] smokers tended to consume more kcal than non-smokers. Few intake differences between smokers and non-smokers were found with respect to fat, protein, carbohydrate, and cholesterol.

Mean vitamin C intake is higher among non-smokers in all age-race-sex categories (Tables 3 and 4) with the difference being most striking among White males. After controlling for age, kcal, PIR, and sex, a negative association remained between smoking and vitamin C consumption for both races.

Univariate analyses of vitamin A intake are inconsistent among age-sex-race categories with non-smokers consuming greater or equal amounts compared to smokers. In multivariate analysis, however, a negative association was found between smoking and vitamin A intake among Whites.

With the exception of older Black males, there is a tendency among all age-race-sex groups for smokers to consume less folate and dietary fiber than non-smokers (Tables 3 and 4). After adjustment, these results were confirmed for dietary fiber in both Blacks and Whites and for folate in Whites only.

Linear Trends in Nutrient and Food Intake

Using multiple linear regression, nutrient intake was examined for a linear relationship between sex-specific quintile of nutrient intake and cigarette consumption, as measured in half packs of cigarettes consumed. No trends were found among Blacks for any food or nutrients after adjusting for age. In Whites, after adjusting for age, negative associations between cigarette consumption and both vitamins C and A were found for males ($\beta = -0.11$ s.e. $= 0.015$, $\beta = -0.06$ s.e. $= 0.012$, respectively) and females ($\beta = -0.14$ s.e. $= 0.014$, $\beta = -0.08$ s.e. $= 0.014$, respectively). Similar negative associations were found for folate among males ($\beta = -0.06$ s.e. $= 0.014$) and females ($\beta = -0.09$ s.e. $= 0.017$) and for dietary fiber among males ($\beta = -0.09$ s.e. $= 0.015$) and females ($\beta = -0.11$ s.e. $= 0.017$). A positive association was found between cigarette consumption and intake of saturated fatty acids for males ($\beta = 0.04$ s.e. $= 0.013$).

For food group intake adjusted for age, a negative linear trend existed between increasing cigarette consumption and intake of any vegetable (males $\beta = -0.10$ s.e. $= 0.028$ and females $\beta = -0.10$ s.e. $= 0.029$), garden vegetables (males $\beta = -0.09$ s.e. $= 0.020$ and females $\beta = -0.10$ s.e. $= 0.027$), any fruit (males $\beta = -0.22$ s.e. $= 0.027$ and females $\beta = -0.34$ s.e. $= 0.030$), fruits or vegetables high in vitamin A (males $\beta = -0.13$ s.e. $= 0.035$ and females $\beta = -0.18$ s.e. $= 0.045$), fruits or vegetables high in vitamin C (males $\beta = -0.21$ s.e. $= 0.031$ and females $\beta = -0.19$ s.e. $= 0.035$), high fiber grains (males $\beta = -0.23$ s.e. $= 0.038$ and females $\beta = -0.15$ s.e. $= 0.048$), low fat milk (males $\beta = -0.17$ s.e. $= 0.025$ and females $\beta = -0.22$ s.e. $= 0.040$), and vitamin and mineral supplements among all Whites. In White males a negative association was found between increasing consumption of cigarettes and salad ($\beta = -0.08$ s.e. $= 0.018$), any cruciferous vegetable ($\beta = -0.09$ s.e. 0.032), and poultry and fish ($\beta = -0.07$ s.e. $= 0.023$).

Discussion

Several limitations of this study should be mentioned. Twenty-four recalls are considered suitable for the study of

TABLE 3—Weighted Mean Nutrient Intake for White Subjects from NHANES II 24-Hour Recall Data by Sex and Smoking Status

	White Men				White Women			
Nutrient	Non-Smoker	95% CI[a]	Smoker	95% CI	Non-Smoker	95% CI	Smoker	95% CI
Ages 19–29 years								
Energy kcal	2947	2790,3105	2997	2868,3127	1667	1614,1720	1679	1603,1755
Protein g	118	112,125	112	107,118	66	64,68	61	57,66
Carbohydrate g	315	300,330	313	297,330	193	186,200	191	182,201
Fat g	122	114,129	118	111,125	68	66,71	66	63,70
Saturated fat g	44	41,47	44	42,47	24	23,25	24	22,25
Cholesterol mg	463	422,505	448	417,479	268	252,284	251	226,275
Vitamin A IU	6000	5368,6631	4529	4155,4903	4582	4156,5007	3810	3142,4477
Vitamin C mg	135	120,150	97	86,108	96	88,104	86	74,98
Folate μg	349	326,372	307	285,330	212	201,223	182	167,198
Fiber g	14.7	13.6,15.7	12.4	11.4,13.4	9.8	9.1,10.5	8.2	7.6,8.8
Ages 30–54 years								
Energy kcal	2444	2340,2548	2569	2493,2645	1535	1488,1582	1591	1542,1641
Protein g	97	93,101	99	96,102	62	60,65	60	58,63
Carbohydrate g	261	248,275	261	254,269	173	168,178	169	163,175
Fat g	102	98,107	108	104,112	63	61,66	68	65,70
Saturated fat g	37	35,39	40	38,41	22	21,23	24	23,25
Cholesterol mg	441	412,469	453	424,482	284	265,302	297	276,318
Vitamin A IU	5796	5112,6480	5926	4854,6997	5567	4726,6409	4204	3676,4732
Vitamin C mg	110	103,117	88	80,95	103	95,111	70	63,76
Folate μg	287	270,305	274	256,292	216	203,229	184	171,198
Fiber g	13.9	13.0,14.7	12.2	11.3,13.1	10.2	9.7,10.6	8.2	7.8,8.6
Ages 55–74 years								
Energy kcal	1998	1955,2042	2010	1942,2078	1360	1329,1391	1390	1336,1445
Protein g	81	79,83	78	75,82	54	52,55	54	51,57
Carbohydrate g	218	212,224	208	199,217	163	158,167	154	146,161
Fat g	82	79,85	85	81,89	54	52,55	57	54,60
Saturated fat g	29	28,30	31	30,33	18	18,19	20	19,21
Cholesterol mg	396	374,418	433	410,456	252	242,262	255	239,272
Vitamin A IU	6320	5824,6816	5990	5278,6702	5426	5000,5851	5225	4575,5874
Vitamin C mg	111	104,118	88	79,97	109	103,115	93	81,105
Folate μg	281	269,293	255	237,272	217	209,224	200	187,213
Fiber g	13.1	12.5,13.7	11.9	11.0,12.9	10.8	10.6,11.1	9.1	8.5,9.8

a) CI = Confidence interval.

group but not individual means.[46] At least one investigator, however, has suggested that the bias due to reporting error inherent in the method may be greater than the difference detected between groups.[47] Given the large size and scope of NHANES II, the 24-hour recall data are likely to be appropriate for the group analysis presented here, but should be viewed with caution because of potential bias. In particular, conclusions regarding nutrients known to have large variability (such as vitamin A) may be subject to question when small groups are examined.[47] Because the number of Blacks sampled in NHANES II is relatively small in comparison to Whites, the results for Blacks cannot be viewed with the same degree of confidence. Finally, the coding of intake of food groups which is based upon mentions from the recalls yields no information regarding portion size and, therefore, is useful but imprecise quantitatively.

Smoking and diet are among the top 10 risk factors thought to cause the majority of disease in the US.[48] Our findings suggest that there are important dietary differences between smokers and non-smokers, all tending toward unhealthy patterns. Surprisingly few differences by current smoking status were found for fat, saturated fat, and cholesterol, dietary variables associated with coronary heart disease. Smokers did, however, consume less poultry/fish and low fat milk and more whole milk and luncheon meats than non-smokers. With respect to cancer prevention, however, there is interest not only in the role of fat, but also in the roles of fruits, vegetables, fiber, vitamin A/carotenoids, vitamin C, and folate. For all these foods or dietary constituents, we found that smokers, in general, consume less than non-smokers, with vitamin C/fruit intake showing the greatest and most consistent difference across all age-sex-race groups.

Our findings with respect to vitamin C and fruits are intriguing for several reasons. First, vitamin C is thought to protect against cancer through reduction functions which serve to both prevent carcinogens from forming and decrease the carcinogenic effects of certain chemical agents and through the enhancement of host resistance.[49] Second, several studies have shown that smokers have lower serum vitamin C levels than non-smokers, and that vitamin C turnover in smokers is higher than in non-smokers.[31,50–52] Several investigators have suggested, therefore, that smokers have increased vitamin C requirements. While mean vitamin C intake data suggest that smokers are meeting their Recommended Dietary Allowance (RDA) for vitamin C (60 mg),[53] this level may not be high enough to counteract the increased vitamin C turnover resulting from their smoking habit. Our data suggest that this very group, which requires more vitamin C, is consuming significantly less. While we do not account for vitamin C intake from supplements, the data show that smokers are less likely to be taking supplements than non-smokers. Third, our vitamin C and fruit results are

TABLE 4—Weighted Mean Nutrient Intake for Black Subjects from NHANES II 24-Hour Recall Data by Sex and Smoking Status

	Black Men				Black Women			
Nutrient	Non-Smoker	95% CI[a]	Smoker	95% CI	Non-Smoker	95% CI	Smoker	95% CI
Ages 19–29 years								
Energy kcal	2603	2247,2958	2796	2348,3244	1715	1576,1855	1731	1573,1889
Protein g	102	88,116	110	76,145	67	62,72	61	55,67
Carbohydrate g	281	241,320	281	253,308	202	176,227	211	187,235
Fat g	110	91,129	114	87,140	71	67,76	67	61,74
Saturated fat g	38	31,44	42	30,54	25	23,28	23	20,26
Cholesterol mg	487	396,578	547	360,735	321	269,372	341	298,383
Vitamin A IU	7973	2410,13536	6530	4701,8358	3923	3306,4539	3987	2829,5145
Vitamin C mg	129	98,160	113	89,138	108	79,136	107	84,130
Folate μg	302	248,356	247	219,274	194	160,227	188	156,220
Fiber g	12.8	10.4,15.2	10.2	8.9,11.5	8.3	7.3,9.3	7.7	6.4,9.0
Ages 30–54 years								
Energy kcal	2006	1784,2228	2369	2140,2598	1391	1268,1514	1328	1239,1418
Protein g	81	70,92	88	81,95	55	50,59	52	47,58
Carbohydrate g	223	204,242	229	209,249	159	143,176	145	129,160
Fat g	82	70,94	93	80,106	57	51,63	56	49,62
Saturated fat g	28	24,33	33	28,39	20	17,22	19	17,21
Cholesterol mg	514	373,655	472	411,533	291	252,331	305	262,349
Vitamin A IU	10204*	6004,14405	4943	3225,6661	4503	3528,5478	6008	4498,7518
Vitamin C mg	133	89,177	104	74,133	96	76,116	71	59,82
Folate μg	264	208,319	213	195,232	173	153,193	155	140,170
Fiber g	12.7	10.3,15.1	9.3	8.4,10.4	7.5	6.8,8.2	7.2	5.9,8.5
Ages 55–74 years								
Energy kcal	1751	1579,1924	1764	1617,1912	1255	1167,1343	1341	1158,1524
Protein g	69	62,76	68	62,75	51	47,54	55	45,64
Carbohydrate g	191	168,214	177	161,193	150	138,162	158	141,176
Fat g	71	63,80	73	64,83	50	45,55	53	45,64
Saturated fat g	25	22,29	27	23,30	17	15,19	18	15,22
Cholesterol mg	378	342,415	420	363,477	273	231,315	261	206,317
Vitamin A IU	6727	4632,8821	5591	4048,7134	7667	5592,9741	4718	2430,7006
Vitamin C mg	98	76,121	79	62,95	114	100,129	105	61,150
Folate μg	216	186,246	241	190,292	190	156,223	175	143,208
Fiber g	11.8	9.8,13.9	10.1	7.8,12.5	9.0	7.9,10.1	9.4	7.9,10.9

a) CI = Confidence interval.
*High value due to 3 individuals with unusually large intakes of liver and/or eggs and/or carrots.

consistent with research suggesting that smoking (via nicotine) decreases the desire for sweet foods and in particular, fruit juices.[54,55] While our results may, in part, be explained by taste preference characteristics due to the metabolic effects of nicotine, this does not resolve the differences found between smokers and non-smokers in their intake of other non-sweet foods.

High intakes of dietary fiber are suspected to be a preventive factor in colon cancer,[56] and the National Cancer Institute recommends that individuals consume 20–30 gm of dietary fiber daily.[57] Past research based on NHANES II data shows that intake of dietary fiber in the US is much lower than this recommendation,[37] and we show that smokers have among the lowest intakes. While tobacco use is not known to be related to bowel cancer, our results suggest that smokers could be at higher risk for this cancer through related dietary lifestyle habits. Furthermore, our results indicate that this lower intake results from lower consumption of several key sources[58] of dietary fiber: vegetables, whole grains, and fruits.

Consistent with past research,[30,32] we found that smokers have, for the most part, lower intakes of vitamin A than non-smokers. While smokers have 25 times the risk of developing lung cancer compared to non-smokers, both low dietary intake and serum vitamin A and/or carotenoids have been shown to increase smoking adjusted lung cancer risk in the range of 1.3 to 2.7 times.[13,14,16,19–25] Furthermore, several studies have shown that low dietary or serum vitamin A/carotenoid levels increase lung cancer risk (range: 1.3–8.1) among smokers, including the heaviest and those who have smoked the longest.[12,16,18,20,22,24,25] Evidence of dietary protection is strongest for carotenoids (from plants) versus retinol sources (from animals).[59] A separate measure of the carotenoid or retinol composition of foods was not available in our database, but our finding that smokers consume fewer fruits or vegetables high in vitamin A, and fewer fruits and vegetables overall, suggest that smokers have lower carotenoid intakes than non-smokers. Stryker, *et al*,[28] have suggested that smokers need to increase their intake of carotenoids to achieve blood levels comparable to their non-smoking counterparts. Our results show that smokers, who may gain some protection from lung cancer by high carotenoid intake, are less likely than non-smokers to be consuming carotenoid-dense foods. Furthermore, former smokers, who enjoy gradually declining lung cancer risk, should be encouraged toward improved nutrition in an effort to possibly accelerate the decline.

A few reports have suggested that folate may be protective for both lung and cervical cancer.[60–62] While our results with respect to intake differences between smokers and non-smokers are less striking for folate than for vitamin C, such differences should be noted in that they fit into a general pattern of lower intake of fruits, vegetables and high fiber grains among smokers.

Questions remain as to whether specific dietary factors such as vitamin A, carotenoids, and vitamin C are the operative cancer protective factors in foods or whether some other yet unknown factor(s) is responsible. Studies have shown that intake of cruciferous vegetables, fruits and vegetables, and especially yellow-orange and green vegetables have protective effects not only for lung cancer, but also for breast, colon, prostate, bladder, oral cavity, stomach and cervical cancers.[56] Our results show that intake of all these foods is significantly lower among smokers. This suggests that smokers may be at higher risk for many cancers, some of which are non-smoking-related. Of particular interest is our finding that for several of the cancer protective food groups, there are modest but consistent trends for intake to decrease as cigarette consumption increases. Thus, the more one tends to smoke, the more one increases cancer risk, the less one tends to consume fruits, vegetables, and high fiber grains. These findings are similar to that of Schoenborn and Benson[8] regarding other lifestyle factors such as sleeping less than six hours per night, skipping breakfast, being less physically active, being sedentary, and drinking heavily.

One possible explanation for lower intakes of nutrients and some foods among smokers may be that smokers compared to non-smokers tend to be greater consumers of alcoholic beverages[5] and, therefore, may replace calories and nutrients from food sources with those from drinks. Because we had no data regarding the contribution of alcohol to either the caloric or carbohydrate content of the 24-hour recalls we analyzed, we were not able to explore this possibility, and past research yields conflicting results.[29,63] Further research is necessary to determine whether alcohol consumption plays a significant role in the lower intake of nutrients and foods found among smokers in this research.

Our findings of lower intakes of vitamin A, vitamin C, fiber, and folate among smokers are correlated in that all these nutrients are common to vegetables, fruits, and grains, foods for which we also found lower consumption among smokers. The task of understanding how nutrients and/or foods protect or promote cancer while considering other risk factors such as smoking is difficult. We have shown that smokers, who are known to be at increased risk for several cancers, have lower intake of nutrients and foods associated with cancer prevention than non-smokers. The evidence is clear that smoking is a habit which leads to or is lead by other lifestyle practices, and we have shown that yet another important health habit is associated with smoking: food intake. This suggests that analyses in future epidemiological research regarding cancer prevention or etiology account for the possible confounding of smoking and diet.

Smoking cessation is the single biggest prevention measure for purposes of cancer prevention and control, particularly with respect to lung cancer.[2] Individuals who smoke are clearly at greatly increased risk from their tobacco use. However, this study suggests that they may be at some slightly additional risk from dietary intake low in critical nutrients. While more research is needed regarding smoking and diet, smokers and ex-smokers may be well advised to increase their intake of fruits, vegetables, and high fiber grains. In the meantime, physicians and the public should be aware that some smokers have low intakes of foods or nutrients associated with decreased cancer risk.

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Depressive Mood, the Single-Parent Home, and Adolescent Cigarette Smoking

LIRIO S. COVEY, PHD, AND DEBBIE TAM

Abstract: The association between depressive mood and cigarette smoking among adolescents was examined within a multivariate model. Subjects were 205 eleventh graders (123 boys and 82 girls) enrolled in a Northeast metropolitan public high school for science-oriented students. Logistic regression analysis showed an independent relation of depressive mood, friends' smoking behavior, and living in a single-parent home with cigarette smoking. Depression scores correlated with the number of cigarettes smoked. These associations suggest that depressive mood and stress may contribute to the onset of smoking. (*Am J Public Health* 1990; 80:1330–1333.)

Introduction

A host of factors have been investigated in numerous attempts to determine the antecedents of cigarette smoking among adolescents. Peer and family smoking behavior have been consistently implicated[1-3]; while less evidence has been demonstrated for the effects of psychological traits, knowledge and attitudes about smoking, school achievement, and demographic characteristics.[1]

A prospective study involving 1,004 subjects found that depressive symptoms reported at ages 15–16 predicted the frequency and duration of cigarette smoking nine years later.[4] This association was not seen for future alcohol or drug use. Another study found higher scores on the Beck Depression Inventory among cigarette smokers than in nonsmokers in males.[5] Both studies limited analysis to simple associations; neither controlled for the effect of other known risk factors. We have found no other attempts to examine the relation between depression and smoking among adolescents. In contrast, a recent paper cited seven studies that examined the relation between depressive disorders and drug and alcohol abuse.[6]

The present study reports our investigation into the relation between depressive symptoms and cigarette smoking among adolescents within the context of a multifactorial model.

Methods

Subjects were 205 11th graders (123 boys and 82 girls) enrolled in a public science high school in a Northeast metropolitan area. This sample included all subjects present on the day our questionnaire was administered in seven non-elective health education classes; they comprised 35 percent of the school's total 11th grade population in the school year 1987–88. Enrollment into this school is competitive and requires passing a qualifying examination. We did not obtain ethnic information from our subjects. School records, however, showed the following ethnic breakdown for the total 11th grade class: 45 percent White, 33 percent Asian, 14 percent Black, 8 percent Hispanic, 0.5 percent American Indian. It is unlikely that our sample deviated markedly from the total 11th grade class in ethnic distribution.

To enhance compliance and candid reporting, we developed a questionnaire that was short (limited to a single, double-sided page), anonymous, and noninclusive of potentially sensitive areas (e.g. grade point average, ethnicity, use of alcohol or drugs). Certain items were not answered by some respondents, accounting for the uneven group totals shown in the accompanying tables.

Subjects were asked whether or not they were regular smokers (at least one cigarette a day), the amount they smoked, the age of smoking onset, and the brand of cigarettes most frequently smoked. We also asked about the smoking behavior of parents, siblings, and friends; estimated parents' annual income; and number of siblings.

We selected 15 of the 20 (CES-D) items[7] to measure level of depressive symptomatology. The feasibility of using this scale with adolescents[8] and that of using a shortened version[9] have been demonstrated. To determine the specificity of an effect of depression, we included two other mental health measures: seven items from an inventory of "worries" used with schoolchildren[10] and from other instruments, assembled seven items assessing level of satisfaction with various life concerns (Appendix). Internal reliability (Cronbach) coefficients based on the total sample were 0.81 for the depression scale, 0.74 for worry, and 0.73 for satisfaction items.

We measured family structure by asking if the subject lived with one or both parents but did not determine whether the custodial parent was separated, divorced or widowed.

Yates corrected chi square analyses and t-tests were used in comparisons of boys vs girls, smokers vs nonsmokers, and highly exposed vs less exposed smokers. Multiple logistic regression[11] using a model that included all the study variables was employed to test the independence of observed associations with cigarette smoking. Quantitative measures were dichotomized at their median values.

Results

We found an overall similarity between boys and girls on most of the psychosocial and smoking variables we studied (Table 1). The mean age of the sample was 16.1 (s.d. = 1.7) years. They were predominantly middle class (about 50 percent estimated their family income to be \$25,000–\$50,000), and living in two-parent households (73.6 percent). Only two children were not living with either parent.

The group prevalence of regular smoking (at least one cigarette daily) was 26.8 percent. It was higher in boys, who also reported smoking more cigarettes, than in girls. Girls reported a lower smoking rate among their mothers than their fathers, a difference that was not seen among the boys (Table 1).

From the Department of Psychiatry, College of Physicians and Surgeons, Columbia University (Covey) and the College of Arts and Sciences, New York University (Tam). Address reprint requests to Lirio S. Covey, PhD, Associate Research Scientist, Department of Psychiatry, College of Physicians and Surgeons, Columbia University, 722 West 168th Street, Box 116, New York, NY 10032. This paper submitted to the Journal November 30, 1989, was revised and accepted for publication March 21, 1990.

TABLE 1—Psychosocial and Smoking Characteristics of 205 11th Graders by Sex

Characteristics	Boys N = 123 (%)	Girls N = 82 (%)	Total N = 205 (%)
Age≤16 years	85 (70.2)	57 (69.5)	142 (70.0)
Mean age (s.d.)	16.0 (2.1)	16.3 (0.6)	16.1 (1.7)
Family income			
<$25,000	17 (16.3)	8 (12.1)	25 (14.7)
$25,000–$49,999	49 (47.1)	36 (54.5)	85 (50.0)
$50,000>	38 (36.5)	22 (33.3)	60 (35.3)
Not the only child	96 (78.0)	63 (76.8)	159 (77.6)
Living with both parents	93 (76.2)	57 (69.5)	150 (73.6)
Regular smokers	34 (27.6)	21 (25.6)	55 (26.8)
Frequency of smoking			
1 cigarette/day	14 (41.2)	12 (60.0)	26 (48.1)
About 10–20	10 (29.4)	6 (30.0)	16 (29.6)
More than 20	10 (29.4)	2 (10.0)	12 (22.2)
Began smoking <15 yrs	19 (55.9)	12 (57.1)	31 (56.3)
Mean age began (s.d.)	14.1 (1.4)	15.5 (6.1)	14.7 (4.0)
Smoke brand with ≥1 mg nicotine	26 (76.5)	10 (52.4)	36 (66.3)
Friends who smoke			
Most or All	23 (22.1)	13 (17.6)	36 (20.3)
Some	49 (47.1)	34 (45.9)	83 (46.6)
None	32 (30.9)	27 (36.5)	59 (33.1)
Mother smokes:	26 (27.1)	9 (11.1)	35 (20.7)
Father smokes:	25 (26.6)	21 (25.9)	46 (27.5)
Siblings smoke:	7 (8.4)	8 (9.9)	15 (10.8)

Note: Percents shown are based on group totals that do not include missing answers.

Mean scores for each of the three mental health variables, i.e. depression, worry, and satisfaction, did not differ by sex. Depression was positively correlated with worry, $r = .30$, (205) $p < .001$ and negatively correlated with satisfaction, $r = -.36$, (205) $p < .001$. Satisfaction and worry scores were negatively correlated, $r = -.16$, (205) $p < .02$.

Rates of cigarette smoking by selected factors are shown in Table 2. The most striking association with cigarette smoking was observed for peer smoking behavior. Smokers were also more depressed and worried, and less satisfied, than were nonsmokers (Table 3).

TABLE 2—Prevalence of Regular Smoking by Selected Characteristics in a Sample of 205 11th Graders

Characteristics	% Smokers	Total N
Age		
≤16 years	23.7	142
>16 years	37.7	61
Family income		
<$25,000	28.0	25
$25,000–$49,999	28.2	85
$50,000>	28.3	60
Only child		
Yes	32.6	46
No	25.2	159
Live with both parents		
Yes	22.7	150
No	37.0	54
Father smokes		
Yes	39.1	46
No	29.8	121
Mother smokes		
Yes	45.7	35
No	29.1	134
Siblings smoke		
Yes	53.3	15
No	27.6	123
Friends smoke		
All	100.0	6
Most	80.0	30
Some	24.1	83
None	8.5	59

Note: Missing answers account for uneven group totals.

Multiple regression analysis (Table 4) showed that three variables were independently related with smoking status. As in the univariate analyses, the strongest relationship was observed for friends' smoking behavior (odds ratio = 7.7). An excess risk for smoking was also seen for respondents who scored above the group median (11>) on the shortened CES-D scale (odds ratio = 3.9), and for those who lived in a single-parent home (odds ratio = 1.9). Only children showed some risk for being a smoker but the confidence interval of the odds ratio (1.6) included 1. No interaction term was associated with smoking.

Smokers who smoked a pack or more a day reported a depression score of 22.7 compared to 15.4 for smokers of fewer cigarettes (difference = 7.3; 95% CL = 2.5, 12.1). No relation between age began smoking or nicotine yield of brand smoked and depressive symptoms was observed.

Discussion

Our sample was drawn from a metropolitan high school for talented, high-achieving students. We can only speculate about the extent to which the competitive and demanding climate of their school environment engendered the preva-

TABLE 3—Mean (standard deviation) Scores of Regular Cigarette Smokers (n = 150) and Nonsmokers (n = 55) in a Sample of 205 11th Graders on Scales Measuring Depression, Worry, and Satisfaction

	Nonsmokers		Smokers		Difference (95% CI)
	Mean	(SD)	Mean	(SD)	
Depression	10.7	(6.4)	17.5	(7.9)	6.8 (4.7, 8.9)
Worry	8.9	(2.5)	9.8	(2.9)	0.9 (0.1, 1.7)
Satisfaction	28.3	(6.7)	25.7	(7.3)	2.6 (0.5, 4.7)

TABLE 4—Multiple Logistic Regression of Smoking Status, Nonsmokers (0) vs Smokers (1) on Selected Characteristics of 205 11th Graders

Independent Variables	Adjusted Odds Ratio	95% CI
Friends smoke (0 = None, Some; 1 = Most, All)	7.70	3.15, 18.8
Depression (0 =<11; 1 = 11>)	3.86	2.56, 5.82
Family structure (0 = 2 parents; 1 = 1 parent)	1.92	1.01, 3.71
Only child (0 = No; 1 = Yes)	1.62	.53, 4.95
Siblings smoke (0 = No; 1 = Yes)	1.13	.50, 2.59
Family income (0 = <$25,000; 1 = $25,000>)	.91	.48, 1.73
Age (0 = <16; 1 = 16>)	.90	.48, 1.68
Father smokes (0 = No; 1 = Yes)	.91	.44, 1.86
Mother smokes (0 = No; 1 = Yes)	.72	.33, 1.57
Worry (0 = <9; 1 = 9>)	.68	.37, 1.25
Satisfaction (0 = <28; 1 = 28>)	.57	.29, 1.16

Likelihood ratio $\chi^2 = 41.84$, df = 11, $p < .001$

lence of depressive symptoms in our sample or the higher rate of smoking we observed (27 percent) compared to the rate seen in similar aged national samples (18 percent).[12] While these special characteristics of our subjects may limit the generalizability of our findings, they do not reduce their validity.

The strong and independent relationship between depressive symptoms and cigarette smoking confirms previous results obtained prospectively[4] and in a cross-sectional design.[8] The correlational nature of our study does not permit us to infer a causal relation or the direction of a causal sequence involving cigarettes and depression. However, neither a direct nor indirect effect from cigarette smoking to depression is likely. There is no evidence that nicotine, the major pharmacological agent in tobacco, causes depression,[13] and our data did not show more depression among smokers of high nicotine-content tobacco. Also, we did not observe that having more friends who smoked (which may reflect membership in a minority, deviant group) is correlated with higher depression. On the other hand, Kandel's observation[4] based on longitudinal data that cigarette smoking in adulthood is a sequela of depressive symptoms in adolescence supports an antecedent, if not causal, relation of depressive affect with cigarette smoking. An association between clinical depression and smoking status has been found in adult samples.[14,15] It remains to be seen whether such an association exists among adolescents as well since the CES-D, while potentially useful as a screening instrument, does not diagnose clinical depression.[16]

The rate of smoking among subjects living in single-parent homes was higher than the smoking rate in other subjects (37 vs 22 percent, respectively), an association that was independent of peer smoking behavior or depression rating. This finding agrees with an earlier observation that disruptive family events are directly associated with adolescent cigarette smoking.[17] In accord with Wills'[18] coping model of smoking in adolescence, cigarette smoking may be an attempt to mitigate the stresses brought on by parents' marital break-up. This explanation is compatible with the notion that parents' divorce may exacerbate the trend towards maladaptive coping among adolescents.[19] We offer these speculations with caution, however, since we obtained only limited information regarding the circumstances surrounding parents' separation.

Our results suggest that merely teaching youngsters how to "just say no" to their peers may not be enough. Depressive mood and parental separation appear to be additional risk factors contributing to the prevalence of smoking among adolescents. Multicomponent prevention approaches that will also address adolescents' emotional needs and the strains engendered by the home situation are indicated.

APPENDIX

Depression Items (15 from the CES-D Scale)

Below is a list of the ways you might have felt or behaved. Please indicate your answer.

a) Rarely or none of the time (less than 1 day)
b) Some or a little of the time (1–2 days)
c) Occasionally or a moderate amount of time (3–4 days)
d) Most or all of the time (5–7 days)

During the past week:

A. I was bothered by things that usually don't bother me
B. I did not feel like eating; my appetite was poor.
C. I felt that I could not shake off the blues.
D. I felt that I was just as good as other people.
E. I felt depressed.
F. I felt hopeful about the future.
G. I thought my life had been a failure.
H. My sleep was restless.
I. I was happy.
J. I talked less than usual.
K. I felt lonely.
L. People were unfriendly.
M. I had crying spells.
N. I felt that people dislike me.
O. I could not get "going".

Worry Items

Almost everyone at one time or another is worried about something. Different people worry about different things. Read and check appropriately if it worries you.

	A lot	A little	Hardly ever
1. Not getting along with your parents	—	—	—
2. Not being popular	—	—	—
3. Feeling left out of things	—	—	—
4. Being made fun of	—	—	—
5. Being looked down upon	—	—	—

Satisfaction Items

For each area of life, choose the letter that corresponds with the level of satisfaction you get from that area.

A) a very great deal
B) a great deal
C) quite a lot
D) a fair amount
E) a little
F) none

a) city or place you live in ___
b) school work ___
c) hobbies or sports activities ___
d) extracurricular activities ___
e) family life ___
f) your friendships ___
g) your physical condition ___

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Effects of Individually Motivating Smoking Cessation in Male Blue Collar Workers

Kaye H. Kilburn, MD, and Raphael H. Warshaw, BA

Abstract: Adverse demonstrable health effects linked to the individual's smoking were shown to 2,689 American workers to motivate cessation during routine examinations to detect asbestosis. This intervention was evaluated six to 25 months later by a mailed questionnaire and by telephone to non-responders. Results were compared to yearly quit rates of 2.5 percent to 5 percent for 736 workers who were ex-smokers at the initial examination. Of the 504 men who responded by mail, 29.8 percent had quit smoking, 35.9 percent had cut down from a mean of 28 to 13 cigarettes per day, and 34.3 percent were smoking as before. Subsequent follow-up at one year showed that 25.6 percent remained quit, and that 23 percent of those who cut down had quit, for an overall quit rate of 34 percent. Of 101 non-responders contacted by telephone, 17 percent had quit and 53 percent had reduced smoking. In both samples, those who quit were more likely to have had lower alveolar carbon monoxide (CO$_a$) levels, to be older, and to have had asbestosis. Responders by mail were the same age as non-responders but had smoked longer, had higher prevalences of asbestosis, emphysema, chronic bronchitis and higher CO$_a$. Demonstration of the adverse personal effects of smoking appear to have contributed to the quit rates or reduced smoking rates in 65 percent of the responding workers. (*Am J Public Health* 1990; 80:1334–1337.)

Introduction

Smoking cigarettes is the largest preventable cause of illness and premature death in the United States. It is responsible for 30 percent of all cancer deaths.[1] Asbestos is the single most important occupational cause of lung cancer.[2] Occupational asbestos exposure raises the lung cancer risk five times and combined with cigarette smoking increases the risks 50- to 100-fold.[3] Smoking and asbestos are synergistic in producing pulmonary fibrosis and functional impairment as well.[4] Despite the fact that blue collar workers in trades exposed to asbestos usually smoke more than others in the work force,[4] few efforts have been aimed at modifying their smoking.[5] In one reported exception, 65 percent of smokers in a California shipyard had quit smoking, leaving only 18 percent of current workers still smoking.[6] This was attributed to their being made aware of their own health risks of smoking combined with asbestos exposure. Thus, we wondered whether other asbestos exposed workers could be motivated to quit smoking if a physician and other health workers pointed out the adverse effects of smoking on their own body and bodily functions. We compared the effects of this intervention to the yearly quit smoking rates in workers who were ex-smokers at the time of study.

Methods

The men studied had been exposed to asbestos in construction trades and shipbuilding as active or retired members of the International Brotherhood of Boilermakers, Iron Ship Builders, Blacksmiths and Forgers, the United Association of Plumbers and Pipefitters or other unions. All men were White . They had to have been exposed to asbestos at least 15 years earlier and for a duration of at least five years. Ages ranged from mid-thirties to mid-eighties. We screened 4,183 of these men from 20 locations in the United States, from Oregon and Southern California to Florida and North Dakota. There were 19.6 percent who had never smoked, 17.6 percent were ex-smokers, and 62.8 percent were current smokers. The intervention target was 2,627 current smokers who received our follow-up letter. The comparison populations for smoking cessation rates were 134 men who had quit in the preceding year and 736 men classified as ex-smokers by interview backed by alveolar carbon monoxide levels.

All men had completed questionnaires in their union halls concerning respiratory symptoms, onset, intensity and duration of cigarette smoking, occupational exposures and exercise tolerance. They had been given a brief physical examination including measurement of height, weight, pulse rate, blood pressure, auscultation of the chest and presence of clubbing, cyanosis, and edema. Chronic bronchitis was diagnosed from the criteria of cough and sputum production for more than three months of two consecutive years.

Smoking status reported on questionnaire, was verified by measuring the level of carbon monoxide in alveolar gas (CO$_a$) during 20 seconds of breath holding with a fuel cell.[4] Vital capacity, forced expiratory volume in one second and flow rates,[7] were measured by spirometry. Each worker's measurements were expressed as percentage of predicted obtained by regression equations.[8] Total lung capacity was measured by planimetering of posteroanterior and lateral chest x-rays, obtained at full inflation of the lungs for area and volumes calculated.[9]

To motivate quitting, the effects of cigarette smoking including symptoms (particularly cough and sputum), chest x-ray changes, abnormal pulmonary function tests and findings on physical examination were explained by the team of examiners. The physician made clear the contribution of cigarette smoking when signs of asbestosis were found in the lung or pleura on the worker's x-ray[10] and when signs of emphysema, such as hyperlucent lung fields, hyperinflation, bullae and other abnormalities[11] were present. He also related these findings to shortness of breath and limited endurance in climbing hills and stairs and observed that these symptoms and abnormalities would progress if the worker continued to smoke but would stop progressing if he stopped smoking. The extremely high risk of dying of lung cancer in asbestos-exposed workers who smoke was also emphasized. Elevated alveolar carbon monoxide (CO$_a$) levels and reduced spirometric flow curves were linked to smoking as were current or eventual shortness of breath by senior technicians.

Address reprint requests to Kaye H. Kilburn, MD, Ralph Edgington Professor of Medicine, School of Medicine, University of Southern California, 2025 Zonal Avenue, HMR 913, Los Angeles, CA 90033. Mr. Warxhaw is a Research Associate at USC. Dr. Kilburn is also affiliated with Workers Disease Detection Services of San Dimas, CA. This paper, submitted to the Journal May 1, 1989, was revised and accepted for publication March 29, 1990.

During the physical examination a nurse practitioner pointed out to those with diminished breath sounds and wheezes that these were characteristic effects of smoking, which indicated obstruction to airflow which in turn causes shortness of breath. Finally, while assisting workers with occupational, medical and pulmonary histories, interviewers emphasized adverse health effects of smoking including the risks for cancer and impairment of breathing which are enhanced by asbestos exposure.

After data were assembled, a personal letter summarizing the findings was sent to each worker from the physician which enclosed results of his pulmonary function tests with the abnormalities identified. Each letter emphasized the greatly increased risk of lung cancer for asbestos-exposed workers who smoke cigarettes.

From six to 25 months after the asbestos screening and study, a letter was sent to each man who was a current smoker. A brief questionnaire was enclosed querying the worker's recall of the screening and report letter, and his current smoking status (quit, cut down, or continued at the same level). He was asked to report the current number of cigarettes smoked per day and whether he still desired to quit. Decreased smoking was defined by the number of cigarettes per day on re-survey being less than the number reported at the examination.

Letters and questionnaires were mailed to 2,689 current smokers; 50 were returned as not reaching the subject and 12 men responded that they had already quit smoking at the time of the screening, which reduced the base numbers to 2,627.

In the early evening, interviewers questioned by telephone 101 of the non-responders selected at random from the 15 largest study sites across the country. Eleven workers who were not reached by calls placed three times on different evenings were replaced. Interviewers asked the same questions as those in the mailed questionnai[illegible] 101 men recalled the asbestos disease screening [illegible] their being strongly advised to quit smoking. Their responses are compared separately from mail responders.

Fifty of the mail responders selected at random from the 150 who reported quitting and 61 selected at random from the 181 who reported cutting down were asked the same questions in a telephone interview 12 months after the return of the questionnaire. Interviewers did not mention the previous mail questionnaire but, if asked, replied that they were doing an interval follow-up. Data from the mail and telephone questionnaires were merged with the data from the asbestosis and lung disease study[4] and tables were constructed showing means and standard deviations. Stepwise linear regressions were made and a correlation matrix was examined.

Results

The data are based on 504 returned questionnaires, a response rate of 19.2 percent to the mailing, and on the 101 non-responders interviewed by telephone. The mail responders were more likely than the 2,123 non-responders to have smoked longer and to have had asbestosis, emphysema, chronic bronchitis and an elevated CO_a level (Table 1), but were otherwise generally similar to responders. Lapse of time since the original study (six to 25 months) was not associated with differences in response rates or in smoking behavior. The 101 non-responders interviewed by telephone were similar to the mail responders in age (mean of 52 years), mean CO_a (34.8 ppm), and prevalence of asbestosis (33.6 percent), emphysema (15.1 percent), and chronic bronchitis (35.9 percent).

TABLE 1—Comparison of Cigarette Consumption, Occupational Exposure, Pulmonary Disease Prevalence and Mean Pulmonary Function Measurements as Percentage of Predicted (pop) in Men who Responded and Did Not Respond to the Smoking Status Questionnaire

	Responded	Did not Respond
Number	504	2123
	mean sd	mean sd
Age yrs	52.9 ± 10.9	53.7 ± 12.0
CO_a ppm	33.0 ± 17.0	22.1 ± 18.3
Duration Smoking yrs	34.2 ± 12.3	27.0 ± 13.4
Number Cig/day	28.3 ± 14.8	28.2 ± 14.8
Cigarette consumed pack yrs	47.9 ± 30.6	3.97 ± 29.6
Current Cig/day	20.4	—
Duration Welding yrs	15.4	14.7
Employed in Shipyards, as %	48.3	48.5
Duration Asbestos Exp. yrs	26.8 ± 10.5	27.7 ± 11.1
Chronic Bronchitis %	38.1	27.7
Asbestosis pleural %	17.1	17.8
Asbestosis parench. %	38.3	28.0
Asbestosis diagnosed, total %	44.0	35.7
Emphysema by x-ray preval. %	15.5	8.3
Bullae prevalence	9.1	4.7
FVC pop	94.7 ± 16.4	94.6 ± 15.8
FEV_1 pop	89.5 ± 22.5	90.7 ± 20.2
FEF_{25-75} pop	77.1 ± 42.4	81.4 ± 40.5
FEF_{75-85} pop	75.3 ± 48.9	80.4 ± 46.8
TGV pop	114.2 ± 14.3	110.8 ± 13.6
RV/TGV	45.4 ± 11.4	44.0 ± 11.5
FEV_1/FVC	69.0 ± 12.8	71.5 ± 11.6

Significant differences compared to "did not respond" are underlined.

Because we lacked a comparison group, we used as historical controls the reported yearly quit rates for the 736 ex-smokers who had been studied for asbestosis and lung diseases at the same time and under identical circumstances to the cigarette smokers and 134 men who quit smoking during the year preceding the study. These were considered to provide the natural smoking attrition rates, before our intervention. The quit rate reported for the year prior to intervention was 4.7 percent and rates varied from 2.5 percent to 5.5 percent, going back 20 years, to 1969.

Of the mail responders, 29.8 percent (150 of 504) had quit smoking after the intervention or after receiving the letter reminding them of their health status (Table 2). Another 8 (1.6 percent) had quit but resumed by the time of inquiry and thus are recidivist smokers. In addition, 35.9 percent reported they had cut down smoking from a mean of 28 to 13 cigarettes per day. The remaining 32.7 percent reported smoking as before but over 90 percent of them said they were still trying to quit. Thus 65.7 percent of the workers who responded by mail had ceased or reduced smoking after the screening and intervention. In the interval telephone follow-up a year after receipt of the questionnaire, the 50 men who had been classified as quitters by mail said they had quit at that time and 43 (86 percent) had remained off cigarettes; seven (14 percent) were recidivists. If these proportions are applied to the whole group, the 25.6 percent of the mail response quitters were not smoking one year after responding to the mail questionnaire. The one-year telephone interview follow-up of 61 men who indicated in the mail questionnaire they had reduced smoking showed that an additional 14 (23 percent) had quit between inquiries and had remained off cigarettes. Applying these quit rates to the entire 504 responders would

TABLE 2—Comparison of Cigarette Smoking, Occupational Exposure, Pulmonary Diagnoses and Pulmonary Function Status of 504 Responders by Mail by Cigarette Smoking Status

	Smoking Status		
	Quit	Decreased	Unchanged
Number	150	181	165
	mean sd	mean sd	mean sd
Age	53.6 ± 10.9	55.6 ± 11.2	49.0 ± 10.1
CO_a ppm	27.5 ± 16.1	33.9 ± 17.5	36.5
Duration Smoking yrs	34.7 ± 11.8	36.4 ± 13.1	31.6 ± 10.5
Number Cig/day	25.3 ± 14.7	28.5 ± 14.0	30.0 ± 15.4
Pack years	45.6 ± 32.6	51.0 ± 30.8	46.3 ± 28.5
Current Cig/day	0	13.3 ± 8.4	28.8 ± 15.9
Duration Welding yrs	15.5	14.9	15.9
Employed in shipyard, %	53.1	49.0	42.5
Duration Asbestos Exp. yrs	27.5 ± 9.5	27.5 ± 11.7	25.2 ± 9.9
Chronic Bronchitis Hist %	33.1	37.0	44.2
Ashthma History %	17.9	19.3	18.8
Asbestosis pleural %	22.7	22.1	6.0
Asbestosis parenc. %	49.3	47.0	18.0
Bastestosis diag tot. %	54.7	55.2	21.2
Emphysema by x-ray prev %	18.0	13.8	15.2
Bullae preval.	8.0	7.1	13.9
FVC pop	95.6 ± 21.2	92.2 ± 17.1	96.4 ± 16.5
FEV_1 pop	90.1 ± 21.2	87.2 ± 23.8	91.2 ± 22.2
FEF_{25-75} pop	75.7 ± 42.9	73.8 ± 43.5	81.3 ± 40.6
FEF_{75-85} pop	77.0 ± 58.5	72.0 ± 44.9	77.2 ± 43.6
TGV pop	113.8 ± 14.4	114.9 ± 13.7	115.0 ± 14.3
RV/TGV	44.8 ± 11.4	48.2 ± 11.0	43.9 ± 10.9
FEV_1/FVC	68.9 ± 12.5	67.9 ± 13.3	70.4 ± 12.1

Significant differences compared to "unchanged" are underlined.

predict that 171 (34 percent) had quit at 18 to 37 months after the intervention. Those who continued to smoke averaged 15.8 cigarettes per day, practically identical to the 15.4 percent reported when contacted by mail 12 months earlier.

Among the mail responders, men who had quit or reduced their smoking were older than those who continued smoking, and had smoked somewhat fewer cigarettes but had smoked longer than those who continued smoking. Initial CO_a levels were significantly lower in quitters than in those whose smoking was decreased or unchanged. Prevalence of asbestosis was over twice as high in those who quit or cut down as compared to those who continued. The lowest mean pulmonary function values and the most pack years of smoking were in men who decreased their smoking but had not quit; they also had more airways obstruction, more air trapping, and lower vital capacities.

The telephone interviews of 101 non-responders showed that 17 percent had quit and remained non-smokers, another 2 percent were recidivists, 53 percent had reduced cigarettes smoked per day, and 28 percent continued smoking as before (Table 3). In the telephone sample, fewer had quit but more had reduced their smoking when compared to the mail sample; of the 34 men with asbestosis, 12 (35 percent) had quit and 17 (50 percent) had reduced their smoking.

Discussion

Quit rates of 17 percent by telephone interviews of non-responders and 30 percent by mail questionnaire were 3.6 to 6.4 times the quit rate of 4.7 percent in the previous year and much above the 2.5 percent to 5.5 percent in the preceding 20 years. Thus, for this blue collar population, the demonstration of their examination, lung function, and x-ray findings and personal advice about present or future health problems from smoking appear to have motivated many of them to quit or reduce their smoking.

The difference in quit rates between the men who responded by mail (30 percent) and non-responders contacted by telephone (17 percent) was not surprising. A 19 percent response is not unreasonable considering the interval of 6–25 months after previous contact. Those with a more socially acceptable outcome, i.e. quitting, would be expected to reply more frequently. Nevertheless, if one applies the 17 percent quit rates of those reached by telephone to the entire

TABLE 3—Categories of Smoking in 101 Telephone Interview Responders by Diagnosis of Asbestosis and Initial Avelolar Carbon Monoxide Level

	Smoking Status				
	Quit	Reduced	Continued	Recidivist	Total
Age years	53.7	51.6	51.6	53.6	
No asbestosis	5	36	24	1	66
Initial CO_a ppm	17.1 ± 8.9	38.5 ± 14.9	34.3 ± 14.1		
Asbestosis	12	17	4	1	34
Initial CO_a ppm	36.1 ± 13.1	33.6 ± 16.9	28.7 ± 14.4		
Total	17	53	28	2	100

mail non-responder group and adds the 171 who quit in the mail responder groups, 531 men are estimated to have quit. This is 20 percent of the 2,627 men screened, at least twice the quit rate of the previous year even if the annual quit rate of 4.7 percent is doubled to make the time intervals comparable.

If decreasing smoking indicates progress toward quitting, then a major proportion (36 percent to 53 percent) of the men were influenced by the intervention. This is supported by the finding in the telephone follow-up that 23 percent of those who had earlier decreased their smoking, quit 12 months later.

Reliability of self reporting has been a controversial subject. Some investigators,[12] have found 95 percent accuracy for telephone interviews compared to personal interviews except for short-term quitters. This reservation would not apply to the long-term quitter. Furthermore, it should be realized that these men did not receive strong or continuous external pressure to quit. They were not even aware, until follow-up, that the intervention was being evaluated. The high correlation between smoking histories and CO_a in the initial study suggested that misleading self-reports about smoking were infrequent.

It is unlikely that the clear differences between quit rates of less than 5.5 percent per year prior to intervention and the 17 percent to 30 percent of new quitters could be due to bias introduced by either the mailed questionnaire or the telephone inquiry. Follow-ups by telephone to those who had responded by mail confirmed their quitting. Veracity was further strengthened by a reported 3.6 percent recidivist rate.

Our intervention was less intense than that in a group of British civil servants who, in hospital because of symptoms or signs of coronary disease or chronic bronchitis or both and received individual advice, were challenged to "consider their situation" and given support and encouragement during four further visits over a period of 12 months. That study had 63 percent of responders (84 percent of the group) who reported being off cigarettes after one year. The present project was also less intensive than a Mayo clinic study of motivated volunteers, mostly referred by physicians in 1975 and seen annually through 1984. They received eight weekly one-hour sessions including presentations from an ear, nose and throat specialist and from a pulmonary physician and achieved a 22.3 percent quit rate at one year,[14] results similar to ours. Nevertheless, our quit rate of 17 to 30 percent of the current smokers plus 36 percent who decreased by 15 cigarettes per day was a substantial beginning.

The men we studied had not volunteered for smoking cessation but were recruited through their unions because of asbestos exposure. Thus, their concerns were whether they had asbestosis, pleural changes, or pulmonary disease. The objective of our intervention was to motivate them to quit smoking, not to help self-motivated volunteers to quit. It appeared to us that the most important motivating factor was measurement of CO_a. Although CO_a had been used previously as a reliability check on smoking histories,[12] in our study it quickly showed these workers a smoking "effect." The smokers' carbon monoxide level was a "big number," in parts per million, compared to that of a nonsmoker or ex-smoker. After CO_a, the two next most powerful motivators seemed to be the x-ray appearance of asbestosis and airways obstruction on the spirogram. Because each individual was "targeted" with personal information, it was a novel experience with strong impact.

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Epidemiologic Differences between Chlamydia and Gonorrhea

Helen L. Zimmerman, BA, John J. Potterat, BA, Richard L. Dukes, PhD, John B. Muth, MD, MPH,
Helen P. Zimmerman, BS, RN, James S. Fogle, and Christopher I. Pratts, BA

Abstract: To assess the prevalence, demographics, and transmission patterns of genital chlamydia infection, we screened 3,078 patients, and compared identified cases (N = 511) to gonorrhea cases (N = 291) diagnosed in the same setting. Chlamydia cases were younger and more likely to be White than their gonorrhea counterparts. Chlamydia cases were distributed diffusely; geographic overlap between the two diseases was only about 40 percent. Gonoccocal coinfection was noted in <10 percent of patients with chlamydia. Nearly half of men with chlamydia and four-fifths of women were asymptomatic and most cases were identified through screening or contact tracing. Populations at high risk for chlamydia are seemingly different from those for gonorrhea. Differences may be due to control interventions (active for gonorrhea, passive for chlamydia). Chlamydia case reporting and control initiatives are recommended. (*Am J Public Health* 1990; 80:1338–1342.)

Introduction

Chlamydia trachomatis causes an estimated three to four million infections annually in the United States; its morbidity, particularly salpingitis, is substantial.[1,2] Nevertheless, only about half of the states require its notification to public health officers.[3] Although much has been learned about its etiology and clinical aspects in the last two decades, little has been published about its epidemiology. Until recently, reliable and inexpensive screening tests have not been available.

To develop a sense of chlamydia prevalence, patient characteristics, and transmission patterns, we implemented a year-long investigation during 1988. Unlike gonorrhea, chlamydia is not a reportable disease in Colorado. Thus we present epidemiologic and case-finding observations on cases of genital chlamydia infection that were detected in the public medical sector, and we contrast some of these findings with gonorrhea cases ascertained in the same setting. In addition, we were able to obtain comparative data from a local military post, where chlamydia and gonorrhea surveillance information is routinely collected.

Methods

Demographics

The study site is Colorado Springs, a standard metropolitan area (SMA) with a population of about 400,000 (1988 estimate), 65 miles south of Denver. Ethnically, the population is 83.5 percent White, 6 percent Black, 8.2 percent Hispanic, and 2.3 percent other races; it is about two years younger than the national median of 31.8 years (1986 estimate), principally because of the large military population. The military post hosts 60 percent of the 35,000 active duty military personnel in the SMA, 92 percent of whom are men. The post population is 65 percent White, 27 percent Black, 4 percent Hispanic, and 4 percent other races. The public clinics are in the SMA's only health department.

Case Definition

The definitive diagnosis of genital chlamydia infection is based on performance of a moderately difficult and expensive cell culture isolation technique. Because this technique was beyond our means, identification of cases and testing of their sexual partners in both the public and military sectors was done with a less reliable but affordable test (MicroTrak, Syva Laboratories, Palo Alto). Hence all cases in our study are presumptively diagnosed.

MicroTrak is a direct immunofluorescence assay using monoclonal antibodies that react with all 15 known human chlamydia serovars; it was performed according to the manufacturer's instructions. Three test outcomes are possible: positive, suspicious (the presence of fewer than 10 chlamydial elementary bodies), and negative.

Test Sensitivity

Because the performance characteristics of MicroTrak were not evaluated by the "gold standard" (McCoy cell culture) in our clinical settings, test outcomes may differ from published reports; these usually claim that the sensitivity of a single genital attempt to isolate *Chlamydia trachomatis* by the McCoy cell culture is about 90 percent for men and between 70 percent and 80 percent for women,[4] and MicroTrak's sensitivity compared to cell culture is about 70 percent for men and 90 percent for women, with a specificity of about 98 percent.[5] Therefore, MicroTrak's ability to detect genital chlamydia in men approaches two-thirds, a value almost identical to the one for women. We use two-thirds as a proportion to estimate the shortfall in our case-finding outcomes.

Patient Selection

Clients presenting for genital examination at our health department STD, prenatal, or family planning clinics were screened for chlamydia during calendar year 1988. Gonorrhea screening in these clinics is universal, except in the family planning clinic, where the gonorrhea culture is used for diagnosis or clinical suspicion and results in about 17 percent of clients being tested. Budget constraints permitted chlamydia screening of only half of the clients in the STD clinic: 52.5 percent of 1,546 women and 55.6 percent of 1,657 men. Although STD clinicians were encouraged to screen younger women (under age 25), all sexual partners of patients with genitourinary conditions and all patients with recurring genitourinary symptoms, the clinicians were not bound by

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From the El Paso County Health Department, Colorado Springs (HL Zimmerman, Potterat, Muth, HP Zimmerman, Fogle); and the University of Colorado at Colorado Springs (Dukes). Address reprint requests to John Potterat, STD/AIDS Control, El Paso County Health Department, 501 No. Foote Avenue, Colorado Springs, CO 80909. This paper, submitted to the Journal July 31, 1989, was revised and accepted for publication March 13, 1990.

these recommendations. All clients in prenatal (573) and all initial visitors to family planning clinic (772) were screened.

Case-finding

Symptomatic men were asked to identify all sexual partners in the interval from one month before the appearance of their urethral symptoms to the date of their treatment; asymptomatic men and all women were queried about their partners in the 120 days prior to the date of their positive test. In the event that there was only one sexual partner in that interval, these latter patients were asked to identify at least one more sexual partner during the year before their diagnosis. The sexual partners (contacts) were actively sought by encouraging self-referral and through field follow-up. All traced contacts were offered epidemiologic treatment as recommended by the Centers for Disease Control (CDC).[6]

Geographic Distribution

The concentration of cases was mapped; census tracts containing approximately 50 percent of cases were considered Core, those containing the next 30 percent, Adjacent, and the remainder, Peripheral, according to Rothenberg's method.[7]

Data Management

Information was recorded on medical charts, case interview records (CDC form 9.97), and contact investigation reports (CDC form 9.2936). The data were coded and entered, without personal identifiers, onto a VAX 8600 computer (Digital Equipment Co.), and they were analyzed using a standard statistical package (SPSS-X, version 3.1).

Results

Population Selection

Younger clients in STD clinic were likely to be tested for chlamydia; the mean age of those tested was 25.3 years (95% CI = 25.0, 25.6) compared to 26.7 years (95% CI = 26.3, 27.1) for untested clients. No tendency to preferentially test by gender or ethnicity was noted.

There were no demographic or case-finding differences between positive and suspicious cases of chlamydia in men; all are included in this report as probable cases. Clinically, 62 percent of the 45 suspicious cases in men were asymptomatic or minimally symptomatic; 19 presented as sexual partners of known cases and 11 had their infection detected as a consequence of screening. The corresponding percentages in the 198 positive cases were: 41.9 (N = 83), 31.3 (N = 62), and 6.6 (N = 13). Concomitant gonococcal infection was observed in 8.6 percent of the positive cases, but in none of the suspicious cases.

Positive and suspicious chlamydia cases in women differed with respect to age, reason for presentation, and case-finding outcomes. Suspicious cases (N = 38) were more likely to be older, with a mean of 25.5 years (95% CI = 23.5, 27.5) versus a mean of 20.7 for positive cases (95% CI = 20.2, 21.2); to present as a result of perceived genitourinary symptoms (39.5 vs 19.8 percent) rather than as contacts or as screenees, and to have fewer chlamydia infected partners in their sexual environments (19.2 percent vs 35.5); 19.2 percent is close to the background level of chlamydia infection in the sexual partners of women who present in our public clinics. Therefore, suspicious cases in women were deemed possible, rather than probable, cases and excluded from the analysis.

Our final study population consists of the 511 cases of chlamydia classified as probable cases in 1988. There were 243 cases in men (198 positive and 45 suspicious cases), all from the STD clinic, and 268 cases in women: 132 from the STD, 75 from prenatal, and 61 from family planning, clinics.

Age Distribution

Sixty percent of the 243 chlamydia cases in men (mean age = 24.2; 95% CI = 23.5, 25.0) and 84 percent of the 268 cases in women (mean age = 20.7; 95% CI = 20.2, 21.2) are under age 25. Fully half of the women are teenagers, as are 20.7 percent of the men. In contrast, 47.5 percent of the 139 gonorrhea cases in men (mean age = 26.1; 95% CI = 24.8, 27.4) and 70.4 percent of the 152 gonorrhea cases in women are under age 25 (mean = 22.9; 95% CI = 21.9, 23.9). Fifty-six percent of men with gonorrhea are in their twenties, whereas 53.2 percent of women are between ages 17 and 22.

Similar differences are noted for the military cases: 80 percent of the 238 chlamydia cases in men whose age was available (vs 72.4 percent of the 337 gonorrhea cases) and 81.7 percent of the 181 age-available cases in women (vs 68.1 percent of the 91 gonorrhea cases) are under 25 vs over 25 years old. The younger distribution in military cases probably reflects the younger age of such populations. The age distribution, by gender, of the sexual partners of our chlamydia patients was similar to that in cases.

Ethnic Distribution

Representation by Whites is higher in chlamydia than in gonorrhea cases for both men and women (Table 1). A similar pattern is noted for military cases, even though the aggregate venereal burden is overwhelmingly Black (Table 2). The race

TABLE 1—Comparison of Ethnic Distribution of Public Clinic Chlamydia and Gonorrhea Cases, by Gender, Colorado Springs, 1988

	Ethnicity			
Cases	White	Black	Hispanic/Other	Total
Men	N (%)	N (%)	N (%)	N (%)
Chlamydia†	150 (61.7)	60 (24.7)	33 (13.6)	243 (100)
Gonorrhea	56 (40.3)	73 (52.5)	10 (7.2)	139 (100)
Women				
Chlamydia‡	161 (60.1)	44 (16.4)	63 (23.5)	268 (100)
Gonorrhea	71 (46.7)	67 (44.1)	14 (9.2)	152 (100)
Total	438	244	120	802

†χ^2 = 30.41; DF = 2; $p < 0.00001$ for the ethnic difference between chlamydia and gonorrhea in men.
‡χ^2 = 42.03; DF = 2; $p < 0.00001$ for the ethnic difference in women.

TABLE 2—Comparison of Ethnic Distribution of Military Chlamydia and Gonorrhea Cases, by Gender, Colorado Spring, 1988

	Ethnicity			
Cases	White	Black	Hispanic/Other	Total
Men	N (%)	N (%)	N (%)	N (%)
Chlamydia†	92 (40)	120 (52.2)	18 (7.8)	230 (100)*
Gonorrhea	52 (15.4)	270 (80.1)	15 (4.5)	337 (100)
Women				
Chlamydia‡	85 (46.7)	82 (45.1)	15 (8.2)	182 (100)**
Gonorrhea	29 (31.9)	51 (56)	11 (12.1)	91 (100)
Total	258	523	59	840

†$\chi^2 = 50.71$; DF = 2; $p < 0.0001$ for the ethnic difference between chlamydia and gonorrhea in men.

‡$\chi^2 = 5.14$; DF = 2; $p < 0.1$.

*Data not available for 20 cases.

**Data not available for 15 cases.

distribution, by gender, of the sexual partners to our chlamydia patients was similar to that of cases.

Geographic Concentration

In the Colorado Springs SMA, the observed chlamydia cases are more diffusely distributed than gonorrhea cases, and the overall extent of geographic overlap between the two diseases is approximately 40 percent. These differences are mapped in Figures 1 and 2. Whereas 50.5 percent of our overall chlamydia cases are observed in 11 Core census tracts (16.2 percent of the SMA's 68 census tracts), nine (11.8 percent) census tracts comprise the gonorrhea Core (49 percent of cases). Only six of the 14 combined chlamydia and gonorrhea Core census tracts (42.9 percent) overlapped. The Adjacent distribution of chlamydia was more diffuse: 29 percent of chlamydia cases occupied 20 census tracts, versus 29.3 percent of gonorrhea cases which occupied 14 census tracts. Overlap occurred in only nine (36 percent) of the 25 combined chlamydia and gonorrhea Adjacent tracts.

The geographic distribution of the sexual partners of chlamydia patients was similar, except that fully 10 percent of contacts emanated from a single military census tract (where *cases* were not enrolled for this study). Exactly two-thirds of all sexual partners resided in Core/Adjacent census tracts, as opposed to about 80 percent of cases. However, the geographic distribution of infected contacts mirrored that of cases: 47.6 percent resided in the Core, and 37.8 percent in the Adjacent, census tracts.

Clinical Findings

One-half (57 percent) of the chlamydia cases in men presented because of symptoms (54 percent had frank urethritis). The remainder were identified through contact tracing (33 percent) or screening (10 percent); these men were essentially asymptomatic.

Chlamydia cases in women were identified through contact-tracing (20.5 percent) or through screening (59.7 percent); only about one-fifth presented spontaneously for examination.

Of the 2,437 public clinic clients tested simultaneously for both infections, 419 had chlamydia, 93 gonococci, and 41 both organisms. Thus 9 percent of chlamydia patients, and 31 percent of gonorrhea patients, were infected with both organisms.

Case-Interviewing

About four-fifths of chlamydia cases were interviewed for sexual partner information (Table 3). No case named a sexual partner of the same gender. The case-interview yield was similar to that obtained in gonorrhea cases.

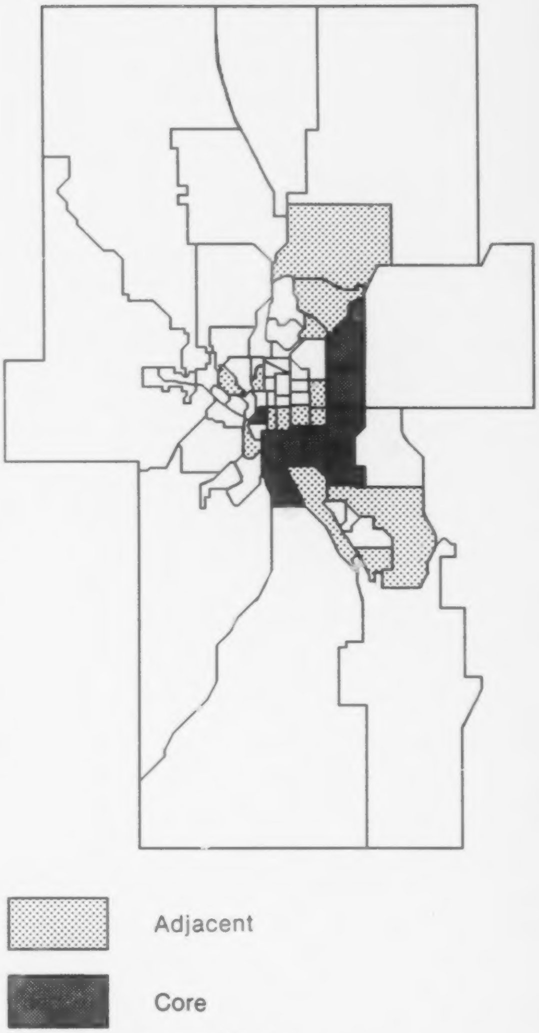

FIGURE 1—Gonorrhea Case Distribution
Public clinic cases (N = 291) in Colorado Springs in 1988: Census tract map showing core and adjacent areas (See Methods)

Had perfect tests been available to use on all examined contacts, the proportion infected would have been roughly the same for contacts to each disease, with chlamydia contacts exhibiting a slightly higher rate of infection in both men and women.

No ethnic differences were noted between chlamydia cases and their contacts.

Case-finding Efficacy

A total of 94 sexual partners had their infection newly identified as a result of case-interviewing and contact tracing (40 female and 54 male contacts). This result produces a 0.22 new case yield (94/419 interviews). Had perfect tests been available, an estimated 141 new cases would have been

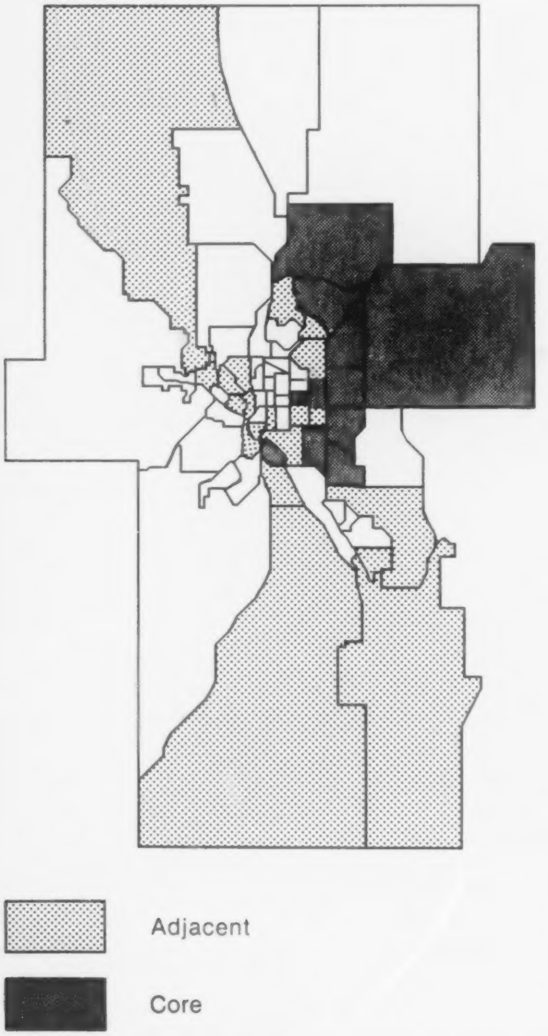

FIGURE 2—Chlamydia Case Distribution
Public clinic cases (N = 511) in Colorado Springs in 1988: Census tract map showing core and adjacent areas (See Methods)

identified, for a yield of 0.34. Thus, we estimate that public health workers need to interview three chlamydia cases and seek their sexual partners (about five), to remove one previously untreated, infected person from the disease pool.

Self-Referral Behavior

Of the 326 named sexual partners who newly were eligible for examination and were found, only 102 (31 percent) were successfully referred by the original chlamydia patient; the other 224 sexual partners required active intervention on our part to assure presentation. Of the 94 newly identified infected contacts, 34 (36 percent) were partner-referred. Black men were least likely to be partner-referred (17 percent), White men most likely (40 percent); Hispanic/Other men, 24 percent. Women were equally likely to be partner-referred (about 22 percent), irrespective of ethnicity.

Discussion

Although ours is not a population-based study of chlamydia prevalence, we found demographic differences between public clinic patients with chlamydia and those with gonorrhea, yet these two diseases infect columnar epithelial tissue and share the same mode of transmission. In Colorado Springs during 1988, public clinic cases of probable genital chlamydia and gonorrhea infection differed with respect to age, ethnicity, and domicile; clinically, little overlap between the two diseases was noted. The same is true for their sexual partners.

Sources of bias in the present study include testing and patient selection. Although neither the sensitivity nor specificity of the MicroTrak test was established locally, it is noteworthy that the estimated proportions of infected persons per sexual partner examined are nearly identical to those obtained in Sweden.[10] If the Swedish data are adjusted for the relative insensitivity of the McCoy cell culture test used in that study, we arrive at 0.84 (estimated) infections identified per Swedish woman contact examined and 0.61 per Swedish man contact, compared to 0.87 and 0.56 (estimated) respectively in our study (Table 3).

The tendency to test younger clients in STD clinic probably affected the age distribution of chlamydia cases. This bias is not likely to be substantial: age differences were also noted in the military cases we reviewed and these were not differentially selected. Bias stemming from incomplete testing for gonorrhea in the family planning clinic probably is minimal: characteristics of gonorrhea cases in this study closely match those from a previous report.[11] The insensitivity of chlamydia testing in asymptomatic men would create bias only in the case of differential insensitivity, which is not likely.

The typical chlamydia case tended to have infection detected as a consequence of screening or contact tracing, rather than via spontaneous presentation with symptoms. It is possible that asymptomaticity is one clue to observed differences in the community expression of these two diseases. In our experience men were at least twice as likely to be asymptomatic with genital chlamydia than with gonorrhea,[8] a ratio similar to that reported in Sweden.[10]

It is possible that sustained, vigorous gonorrhea control efforts on our part since the mid-1970s[12,13] influenced the observed distribution of chlamydia cases in our community. In Colorado Springs this population is principally White. As our interventions helped to decrease the gonorrhea burden from a high of about 2,000 cases in 1977 to about 1,000 cases in 1988, the proportion of cases accounted for by non-Whites increased from 39 percent in 1977 to 67 percent in 1988. This phenomenon may also explain the modest proportion of clinical overlap between the two diseases since gonorrhea is relatively well controlled in Whites locally.

We have demonstrated previously that gonorrhea is neither randomly nor evenly distributed in our sexually active populations[11] and that the basic setting for sexual aggregation and for sexual choices is the neighborhood. It is possible that in neighborhoods where the gonorrhea burden is high, the chlamydia burden may be lower than expected because of frequent exposure to antibiotic treatment on the part of gonorrhea cases and their sexual partners, especially since the introduction of dual treatment for gonorrhea and chla-

TABLE 3—Chlamydia Case—Interview Outcomes in 1988, Compared with Previously Reported Gonorrhea Case-Interviews in Colorado Springs, by Gender

Cases						
Disease	Interviewed	Contacts Elicited	Contacts/ Case	Examined (%)	Infected‡ (Observed)	Infected** (Expected)
Men				Contacts (Women)		
Chlamydia	190	321	1.7	176 (55)	98 (.58)	147 (.87)
Gonorrhea*	394	827	2.1	533 (64)	381 (.71)	423 (.79)
Women				Contacts (Men)		
Chlamydia	229	379	1.7	254 (67)	94 (.37)	141 (.56)
Gonorrhea†	578	1492	2.6	1042 (70)	393 (.38)	491 (.47)

*Data from reference 8

†Data from reference 9

‡Per contact examined

**Expected if laboratory tests were perfect. For chlamydia we use two-thirds as the observed sensitivity for men and women (see Methods) and for gonorrhea, 90% sensitivity for men and 80 percent for women.

The infection rate for women contacts to men with chlamydia may be underestimated because 21 of the 176 "examined" women were treated by their private physician without being tested.

mydia in the early 1980s.[6] Conversely, in neighborhoods where gonorrhea is an infrequent event, exposure to chlamydia may result in successful infection more often. In addition, entrenchment and perpetuation of chlamydia may be facilitated by its high rate of asymptomaticity. Age differences between chlamydia and gonorrhea patients may be accounted for by a tendency on the part of teenage women with chlamydia to select sexual partners close to their own age.

More research is needed to elucidate the factors that contribute to differences in the community distribution of these two closely related diseases. Meanwhile it is clear that public health intervention is crucial to the control of chlamydia infection in Colorado Springs. The disease is frequently asymptomatic, occurs often in populations that we are not encountering routinely in our standard STD outreach efforts (syphilis, gonorrhea, and human immunodeficiency virus infections), and the chlamydia-prone populations are only moderately successful at referring their sexual partners. Our data support a recent appeal to implement a "program of chlamydia control that is based on accurate diagnosis, contact tracing, appropriate therapy, and educational efforts to reduce risky behavior . . ."[14] Our first step is to encourage the reporting of chlamydia cases to public health officers.

ACKNOWLEDGMENT

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A Small Area Simulation Approach to Determining Excess Variation in Dental Procedure Rates

PAULA DIEHR, PHD, AND DAVID GREMBOWSKI, PHD

Abstract: All small area analyses need to compare the observed variability in rates to that expected by chance alone, but the expected variability is usually not known. This paper uses patient-level data for five dental procedures to simulate the distributions of the summary statistics that are usually generated in such studies. These statistics are found to vary greatly even under the "null hypothesis" that all dentists are using procedures at the same rates. The simulated dentist rates are compared to observed rates obtained in a different study. These findings illustrate problems that can occur in small area analysis studies, and emphasize the importance of using statistical techniques that are appropriate for the data that are to be analyzed. Investigators should make every effort to obtain patient-level data, or at least to understand the underlying distribution of the number of procedures per patient, to avoid mistaking significant deviations from an incorrect model as evidence for significant variation among small areas. (*Am J Public Health* 1990; 80:1343–1348.)

Introduction

Small area analysis is a popular methodology in health services research. A typical study might calculate the utilization rate for a service in several small areas, compare the largest rate to the smallest, note that the difference is large, and attempt to explain the high variability as a function of service availability, physician uncertainty, and other variables of interest. The statistical methods used in such studies have been questioned by an author of this paper.[1]

A recent paper using computer simulation to examine the statistical properties of the most frequently used descriptive statistics[2] demonstrated that all of the usual descriptive statistics could be deceptive, showing large apparent variation when there was no more variation than would be expected by chance alone. A simple 2 × k chi-square test (classifying people in each of k communities into two cells, by whether they had or did not have the procedure) was an appropriate test for excess variability as long as each person could have the procedure at most once. However, the chi-square test could be deceptive; results would be "statistically significant" too often if an individual could have the procedure more than once. Simulation studies were recommended for these situations; this requires that the distribution of the number of procedures per person be known. Finally, it was pointed out that studies very similar to small area analyses were being conducted in other situations: e.g., a "small area" might be a hospital or a dental practice. This paper applies the small area simulation approach to a set of dental data, to determine if there is more variation in procedure rates among dentists than would be expected by chance alone.

Grembowski, *et al*,[3] calculated procedure rates for 200 dentists in general practice in four urban counties of Washington State from 1984 and 1985 dental claims of members of the Washington Education Association (WEA) and their dependents. The number of patients seen at least once in the two-year period ranged from 75 to 300 for the 200 dentists, as shown in Table 1.

TABLE 1—Number of Washington Education Association Patients Seen by each Dentist in Two Years

N of Patients Seen	N of Dentists
75–99	94
100–124	54
125–149	19
150–199	24
200–300	9

Procedure rates for each dentist were calculated by dividing the total number of procedures performed in a two-year period by the number of people seen at least once by the dentist during that period. Results were age-adjusted. Table 2 shows the substantial variability that was found among dentists: rates of oral examinations ranged from 10 per 1,000 patients to 3,543 per 1,000 patients. The coefficient of variation (the standard deviation divided by the mean) was 0.289. This seems like substantial variation. However, a good deal of variation can be expected by chance alone, even if the underlying rate of examinations per person is the same in each practice.[2] If "exams" were a binary variable, which a person could have once or never, it would be appropriate to use a 2 × 200 chi-square test to examine the null hypothesis that the probability of having an oral exam was the same in all practices.[2] However, it is obvious that many people have more than one examination in two years, which means that the chi-square test and the usual small area statistics are either inappropriate or have unknown distributions under the null hypothesis.

If the number of oral exams for each person were known, techniques could be developed to study whether there was excess variation. The data could be recoded so that each

TABLE 2—Observed Procedure Rates in 200 Dental Practices (number of procedures per 1,000 patients)

Category	Mean	SD	Min	Max	EQ	CV
Oral Exams	1617	467	10*	3543*	354.3*	.289*
Fillings	889	373	188*	2379	12.65*	.420*
Extractions	119	83	0	529	—	.698*
Root Canals	44	40	0	228	—	.909*
Scalings	23	47	0	345*	—	2.044*

EQ = Maximum/Minimum.
CV = coefficient of variation.
*Minimum lower than 5th percentile of null distribution, or maximum higher than 95th percentile (see Tables 3 and 4).

Address reprint requests to Paula Diehr, PhD, Professor, Department of Biostatistics, SC-32, School of Public Health and Community Medicine, University of Washington, Seattle, WA 98195. Dr. Grembowski is Assistant Professor of Health Services and Dental Public Health Sciences at the University of Washington. This paper, submitted to the Journal September 12, 1989, was revised and accepted for publication March 8, 1990.

person was coded 1 if he had one or more exams, and 0 otherwise, and the chi-square statistic could be used; or, one could assume that the number of exams per person was approximately normally distributed, and perform an analysis of variance to study whether the average number of exams per person were different among the dentists. Unfortunately, the only data available for the Washington study were the rates per dentist of the various procedures. No patient level data were available. For these reasons, we attempted a simulation approach, discussed elsewhere in detail.[2] The goal is to determine what the distribution of procedures among the 200 practices would have been under the null hypothesis that no dentists were "high users" but, on the contrary, that all dentists performed the same number of procedures per person, on average. The simulation method attempts to determine how much variation among practices would be expected by chance alone. This is then used as a benchmark to determine whether the observed variation in Table 2 is important.

Methods

Data

A different data source was used to obtain information about the distribution of procedures at the patient level. Annual dental utilization data were obtained for 12,897 insureds from a random sample of adults and children covered by Pennsylvania Blue Shield dental insurance in 1980. These data have been analyzed elsewhere.[4,5] The 5,912 people with no dental visits in 1980 were eliminated from the data set. Appendix 1 contains frequency distributions for five dental procedures—oral exams, fillings, simple extractions, root canals, and full-mouth periodontal scalings. (For example, 2,764 patients had no oral exams, 2,717 had one exam, and six people had 12 oral exams in one year.)

Fitting the Pennsylvania Data

We first attempted to fit the Pennsylvania data to the Poisson and negative binomial distributions.[6] If a known distribution was appropriate, we could use it to generate random data for the simulation. Clearly, none of the data are normally distributed because of the long tails and the high number of observations clustered at zero. The Poisson distribution is not likely to fit the data because a Poisson process assumes that procedures occur independently, which is probably not the case for dental procedures. The negative binomial distribution assumes that each person uses procedures with a Poisson distribution, but that each person has a different Poisson parameter, distributed according to a gamma distribution.[6] This distribution was considered because it has longer tails than the Poisson distribution.

We estimated the parameters of the Poisson and negative binomial distributions using the maximum likelihood procedure.[6] This yielded expected numbers which were compared to the observed numbers with the chi-square goodness of fit test.[7] Neither distribution provided an acceptable fit for the number of oral exams or for fillings. The negative binomial distribution did provide an adequate model for the number of extractions, root canals, and scalings.

The following simulation work did not assume any mathematical distribution, because there was none that worked for all the data. Rather, data were generated directly from the empirical distributions, as described below. Future work in simulating procedure rates for extractions, root canals, and scalings might use the negative binomial distribution to generate random data, however.

Simulation of Utilization Rates

Our goal was to develop a simulation model to assess the amount of variability that would have occurred among the dental practices by chance alone if the null hypothesis (that all dentists use procedures at the same underlying rate) is true. We generated 100 different simulated sets of procedure rates, in which the null hypothesis was true.

For each of the 200 dentists, a uniform random number between 0 and 1 was generated for each patient in that practice (e.g., for a dentist with 75 patients, 75 random numbers were generated). Each of these numbers was compared to the cumulative distribution in Appendix 1. For example, 2764/6985 or 39.5 percent of the people had no oral exam; 78.46 percent had 0 or 1 exam; 97.7 percent had 0, 1, or 2 exams, etc. For each random number, if it was less than .395, the patient was assigned 0 exams; if between .395 and .7846, the patient was assigned one exam, etc. Thus, on average, the number of exams per patient match the distribution of Appendix 1, but the distributions were somewhat different for each dentist.

After the number of exams were generated for all 75 patients in the first practice, the exams were summed, multiplied by 2 (see discussion section) and divided by 75 to provide a simulated two-year rate of oral exams for that dentist. This was then continued for all 200 dentists, which generated 200 different simulated examination rates. Next, the minimum rate, the maximum rate, the extremal quotient (EQ, ratio of the maximum rate to the minimum), the coefficient of variation, and other statistics explained in the results section were calculated. This process was repeated 100 times to give, for example, the "average" maximum rate, or the 95th percentile of the maximum rate, under the null hypothesis that each practice has the underlying distribution of examinations shown in Appendix 1.

Results

Mean Procedure Rates

Table 2 shows the observed results, and Table 3 shows the simulated results. The first column in Tables 2 and 3 is the estimated mean number of procedures per 1,000 patients. Recalling that the two sets of numbers arise from different data sets, there is fairly good correspondence on the rate of oral exams (1,725 versus 1,617), root canals (55 versus 44), and full mouth scaling (34 versus 23). There is very poor agreement on fillings (2,123 versus 889) and for simple extractions (211 versus 119). This may be due to differences in billing practices between Pennsylvania and Washington Dental Service—the carrier for Washington Educational Association. There are probably also socioeconomic differences between the Washington and the Pennsylvania patients. These differences are considered further below.

TABLE 3—Simulated Procedure Rates (number of procedures per 1,000 patients)

	Rate	Minimum		Maximum		EQ	
	Mean	Mean	5%	Mean	95%	Mean	95%
Oral Exams	1725	1260	1136	2304	2537	1.8	2.1
Fillings	2123	1195	974	3422	3897	2.9	3.4
Extractions	211	30	0	529	667	17.5	30.1
Root Canals	55	0	0	259	342	—	—
Scalings	34	0	0	227	316	—	—

EQ = Maximum/Minimum
Based on 100 iterations per procedure.

Maximum and Minimum

The simulated minimum and maximum are shown in Table 3. For example, on average, the lowest rate of oral exams was 1,260 per 1,000 patients, and the highest rate was 2,304 per 1,000, even though the underlying rate for all dentists was 1,725. Thus, an observed minimum rate near 1,260 or a maximum rate near 2,300 would not be an indication of excess variability among practices. Table 3 also shows the simulated 5th percentile of the minimum and 95th percentile of the maximum. That is, 95 percent of all minimum values are above 1,136, and 95 percent of all maximum values are below 2,537.

The observed minimum rates for oral exams, shown in Table 2, can be thought of as extremely low, since 10 is substantially below 1,136. The same is true for fillings. The minimum for the other procedures is 0 for both the observed and the simulated data. Comparing the maximum values of Table 2 to the 95th percentiles in Table 3, the maximum rates for oral exams and for full mouth scaling are higher than would be expected by chance alone, but the other maxima are not particularly high.

Extremal Quotient (EQ)

Table 3 shows the simulated average value and the 95th percentile of the EQ (maximum divided by minimum). For oral exams, one would expect an EQ of about 1.8, and values above 2.1 (the 95th percentile) would be considered particularly high. For fillings, a ratio of 2.9 is expected, and 3.4 would be considered high For extractions the minimum value was 0 in 10 percent of the simulations, and the EQ could not be computed. For the remaining 90 percent, the expected EQ was 17.5 and the 95th percentile of the EQ was 30.7. The EQ is infinite for root canals and scalings, as the minimum is always zero. In Table 2, the observed EQ could be computed only for oral examinations and fillings. Both of the observed EQs are much larger than the 95th percentile of the EQ distribution, indicating that the null hypothesis that all dentists had the same rates is probably not true. The variability in the rates is due in part to some dentists having much *lower* use of procedures than would be expected.

Coefficient of Variation (CV)

Table 4 shows that the 95th percentile of the simulated CV (standard deviation over mean) for oral exams is 0.115, as compared to an observed value of 0.289 shown in Table 2. The observed variation is thus higher than would have been expected by chance alone. All of the observed CVs are considerably above what would have been expected. Therefore, the null hypothesis probably is not true, and there are some dentists who use procedures at different rates from the others.

TABLE 4—Simulation Results—Other Summary Statistics*

95th percentiles (per 1,000 patients)					
	CV	SCV	199 df Chi sq	1 df Chi sq	F
Oral Exams	.115	8	−441	7.28	1.22
Fillings	.194	35	−1116	27.31	1.13
Extractions	.442	148	1070	13.15	1.17
Root Canals	.914	674	1035	14.35	1.19
Scalings	1.188	1468	1208	18.97	1.20
Degrees of Freedom			199	1	199, inf
Tabled 95th Percentile			233	3.84	1.17

*CV is the coefficient of variation and SCV is an estimate of the systematic component of the variation, defined in the text. Chi-square with 199 degrees of freedom is 2 × k chi-square assuming that each procedure is a person. Chi-square with 1 df is the test for a particular practice (see text). F is the F statistic that arises from a 1-way analysis of variance. Tabled 95th percentiles are taken from tables of the chi-square and F distributions.

Based on 100 iterations per procedure.

Systematic Component of Variation (SCV)

Table 4 shows some other statistics commonly used in small area analysis, which were not calculated for the Washington data. They are included for completeness. The SCV has been proposed by McPherson and Wennberg as a test for excess variability among small areas.[8] The formula is:

$$SCV = (1/\mathrm{k})\ [\Sigma((\mathrm{O}_i - \mathrm{E}_i)^2)/\mathrm{E}_i^2 - \Sigma(1/\mathrm{E}_i)]^*\ 1000.$$

There is an F-test associated with this statistic, but it is not often used. The SCV was derived under the assumption of Poisson rates, which is not appropriate here. The data in Table 4 illustrate further what has been shown elsewhere[2]; that is, that the SCV can take on very large values, even when there is no underlying variation among the small areas. In addition, it becomes larger as the mean rate becomes smaller. Thus, it would be unwise to use the SCV alone as a descriptive statistic in this situation.

Chi-Square

Two chi-square tests have been used in small area analyses. One test is a 2 × 200 chi-square test, which assumes that each person received a procedure at most once, calculates the number not receiving the procedure as "population size minus the number of procedures," and creates a 2 × 200 table. This is clearly inappropriate in our situation, since many people had more than one procedure. In fact, the 95th percentile of this chi-square statistic is negative for oral exams and fillings, because the expected numbers of people who did not have the procedure are negative for some dentists (for example, there are more fillings than patients). A chi-square value greater than 1,000 would be needed to demonstrate "significance" for the last three procedures, as compared to the tabled 95th percentile of the chi-square with 199 degrees of freedom, which is 233.

Another test which is proposed in these situations is a 1 degree of freedom chi-square test of the hypothesis that a particular dentist has a rate significantly different from what would have been expected. The expected number of procedures would be the grand rate (e.g., 1,725 oral exams per 1,000) times the number of patients (e.g., 75, for an expected rate of 129 exams). If the observed rate were, e.g., 155, the chi-square statistic would be $X = (O - E)^2/E = (155 - 129)^2/129 = 5.2$. This is much higher than the usual critical value of 3.84 for the chi-square distribution with 1 degree of freedom. However, Table 4 shows that the 95th percentile of chi-square statistics in this situation is 7.28, meaning that 5.2 is not a particularly large value. The 95th percentiles for the other dental procedures are even larger, ranging from 13 to 27. Thus, the simple chi-square test, to look for outliers, could be extremely deceptive in this situation, indicating significant variation when a practice was operating, on average, like all other practices. Or, put another way, 16 percent of the chi-square statistics would have been greater than 3.84 for oral examinations, giving a Type I error of 16 percent rather than the commonly accepted 5 percent level. The other Type I errors are 47 percent for fillings, 30 percent for extractions, 28 percent for root canals, and 19 percent for scalings. If these data were being used to identify outlying dentists, far more than 5 percent would be incorrectly labeled as over-providers. This discrepancy occurs because the data violate the essential property required for the chi-square test:

that the counts be independent (or, that the distribution of counts per person follow the Poisson distribution).

There is another problem with this chi-square test, which is that it ignores multiple comparisons. If a chi-square test, even using the correct 95th percentile, were performed for all 200 dentists, about 5 percent (10 practices) would be judged to have excessive utilization rates by chance alone. The probability that no dentist was classified incorrectly is $.95^{200}$ = .00004. Thus, the experiment-wise Type I error is almost 1.0. This might be adjusted for by using an alpha level of .05/200 = .00025 instead of alpha = .05. This would require comparing the observed chi-square values to the 99.975th percentile of the distribution (about 17.17 instead of 3.84 in this situation).

Clearly, the combination of these two problems (that the chi-square statistic does not follow the chi-square distribution, and that there are multiple comparisons) will tend to yield extremely high observed chi-square values even in the null case. Users who are not aware of this problem risk labeling small areas as outliers that are not extreme at all. We have seen examples in which a large percent, or even *all* of the small areas being considered, were labeled as outliers based on this test![9]

Analysis of Variance (F-test)

The 95th percentile of the F distribution is given in Table 4. This statistic would have resulted if we had performed an analysis of variance to detect a difference among the 200 dentists in the mean number of procedures per person. Analysis of variance requires having patient-level data, so that the variation among people can be computed. This was not available in the Washington data, but could be computed from the Pennsylvania distributions. The 95th percentile for an F distribution with 199 and infinite degrees of freedom is 1.17 if the null hypothesis is true. This is close to the simulated F values for the dental procedures (1.22, 1.13, 1.17, 1.19, 1.20), showing that analysis of variance would have been appropriate if patient-level data had been available, even though the data in Appendix I are far from normally distributed. (We used only 100 iterations, so the percentiles might be even closer than is shown.) If the data are not available at the individual level, to permit analysis of variance, it will be necessary to perform a simulation study to determine whether there is more variability than would be expected by chance alone.

Adjusting the Pennsylvania Rates

It is unfortunate that some of the average rates were so different in the Washington and Pennsylvania data sets. This suggests that some of the "significant" variation detected could have been caused not by excess variation but because the Washington data were being compared to the wrong underlying distribution. To examine this possibility we standardized the results, so that the average simulated rate would be the same as the average observed rate. Specifically, we multiplied the simulated rates by a constant factor (e.g., for oral exams 1617/1725 = 0.937 so that the simulated mean would be 1617). This has the effect of multiplying most of the values in Tables 3 and 4 by a factor less than one (0.937 for oral exams). (The EQ, CV and F values would be unchanged). The observed data would then be as extreme or even more extreme, if compared to these adjusted tables. All of the observed maxima would be larger than the adjusted 95th percentile except for root canals.

Discussion

The intent of this paper was to demonstrate that small area analysis issues were similar to those in detecting excess variation in other settings, and in particular to illustrate the simulation approach. Both the findings and the methods merit some discussion.

If we assume that the underlying distribution of procedures is about the same in the Washington and Pennsylvania data, the simulation method has demonstrated that there was excess variability for oral examinations and full mouth scaling. The other procedures also showed excess variability, although not as unequivocally.

There are some shortcomings to these results, and also to the methods that we used to obtain them. The Washington population was a fairly homogeneous group of middle-class patients. The Pennsylvania Blue Shield data were from a random sample of all insureds, who varied by social class. The differences between the two study populations may explain why the mean rates for the rehabilitative services in Table 3 were higher than those in Table 2. Also, Pennsylvania had better coverage for fillings than did the Washington contract, which may account for the discrepancy in filling rates between Tables 2 and 3.

The estimated rates from one year of Pennsylvania data were multiplied by 2 to give two-year rates for comparison with the Washington data. This is not quite appropriate, as it assumes that non-users in the first year would not have used any services in the second year. The correct multiplier is between 1.08 and 2.0 (details from authors). The "2" was chosen as an upper bound, to provide as much variability as possible. If 1.08 had been used, all of the observed maxima would have been above the 95th percentile.

No effort was made in either data set to remove people who were not covered for the entire period of time. If a person was insured for only one month of the two-year period, the rate was calculated as though the person had been at risk for the entire 24 months. This would tend to lower the rates, and would lower them more for the Washington data than for the Pennsylvania data, since the time period was longer for Washington. This may explain some of the differences in the two data sets.

We did not model one aspect of the variability—that a different number of patients might have been seen in different two-year periods. Age and sex differences were not incorporated in the simulation, as the Washington data had already been age standardized, and sex is not an important determinant of use in insured populations.[4,5] We showed elsewhere[2] that this was a reasonable way to proceed. Finally, the estimated means and percentiles are not completely accurate, as there were only 100 simulations per procedure.

Once excess variability has been established, it is appropriate to examine reasons for this variability. Grembowski, *et al.*,[10] have shown that structural features of the practice, such as its age and size, explain some of the variation in the rates.

Although not completely satisfactory, the simulation approach has provided support for the finding of excess variability among dentists. Utilization review programs operated by dental insurers are often implemented to identify dentists who have high procedure rates. The insurer may, for example, intensify review of the dentist's claims as a means of reducing the dentist's rate. Our results indicate that insurers need to determine whether excess variability exists for a given procedure before a dentist is labeled as providing "too many" services. Of equal importance, if excess variability is detected, dentists with the lowest rates may be underutilizing appropriate services.[3] Therefore, insurers should target utilization review at both ends of the rate distribution.

The methodology of this study is applicable in other situations. We used the simulation method to examine the characteristics of small area analysis statistics applied to dental practices instead of geographic areas. Dental practices have smaller numbers of patients in each "small area" and some of the procedures are applied multiple times to the same person. The distributions of root canals or scalings, which are rare, are more similar to the distribution of the number of hospital admissions for a particular diagnosis, which is often considered in small area analyses. The main results of this study have to do with the underlying distribution of services, the properties of several descriptive statistics, and various hypothesis testing procedures.

None of the distributions of services in Appendix 1 had a Poisson distribution. This is important because several approaches are now being used in the small area literature which assume that the underlying distribution is Poisson: the 1 df chi-square test and the SCV were derived under these assumptions[8] and some regression approaches model deviations from the Poisson distribution.[11,12] In all of these methods, statistically significant departures from the Poisson distribution are taken as evidence that there is unexplained variability among the small areas. However, as shown here and elsewhere,[2] the underlying distribution at the person level may not have a Poisson distribution, and finding "significant" variation is as likely to be caused by this fact as to represent excess variation among the small areas. Unless there is evidence that the underlying distribution is Poisson (e.g., that a person cannot have more than one of the procedures of interest) this is an inappropriate inference.

Tables 3 and 4 demonstrate that a considerable amount of variability can be expected among small areas by chance alone, and that it is dangerous to "eyeball" descriptive statistics as a test for excess variability. The minimum and maximum rates, and the extremal quotient, could vary considerably under the null hypothesis. The coefficient of variation and the systematic coefficient of variation also vary substantially, and tend to be larger for the less prevalent procedures (i.e., higher for root canals and scaling than for oral exams and fillings). It is common to compare CVs or SCVs across several procedures, and to claim that those with the largest coefficients represent "practitioner uncertainty." If we had done this with the data of Table 4, we would have "found" that there is less certainty about root canals and scaling than about oral exams and fillings, even though under the simulation model all dentists used procedures at exactly the same underlying rates! It is inappropriate to infer differences in "certainty" based only on CVs or SCVs if procedures have different mean rates.

Several tests of the hypothesis that there is underlying variation among the small areas have been developed. One is a 2 × k chi-square test which is appropriate if a person can have the procedure of interest at most once (e.g., death, removal of an organ). This procedure failed for the data of this study, even for procedures such as root canals, where the assumption that few people had more than one per year would have seemed "reasonable." Thus, the assumptions of this test must be examined with care.

The second chi-square test, which purports to test whether an individual dentist or small area is an "outlier" (observed significantly different from expected) was also shown to be inappropriate first, because the data do not meet the underlying assumption of a Poisson distribution, and second, because the problems of multiple comparisons are ignored. It is very important that this test be used correctly, since it singles out particular areas or dentists as outliers, and may have deleterious effects.

A third test which is sometimes proposed is a simple analysis of variance. The simulation results of Table 4 show that it would have been appropriate for these data, if data had been available at the patient level. (Such data may be difficult to obtain, as there must be a unique patient identifier to permit ascribing multiple procedures to a person.) If investigation of patient level data showed that they were inappropriate for analysis of variance, logarithmic or other transformations could be attempted.

The simulation method we have proposed has so far been used only in the null situation. The power of the various procedures to detect true underlying variation has not been studied; it should be studied. It is possible that tests based on the CV, for example, would be more powerful than other types of tests. More methods research is needed in this area.

The major lesson of this paper is that it is important in small area analysis to understand the underlying distribution of services, at the patient level, to avoid mistaking misspecification of the underlying model for statistically significant excess variability among small areas.

DISCLAIMER

Interpretations of the data are the authors' own and do not necessarily represent those of the National Center for Health Services Research and Health Care Technology Assessment, Washington Dental Service, or the Washington Education Association.

APPENDIX 1

Number of Procedures per Person in Pennsylvania Data*

Number of Procedures	Procedure				
	Exams	Fillings	Root Canals	Full Mouth Scaling	Extractions
0.	2764	4020	6839	3722	6446
1.	2717	1239	120	43	417
2.	1349	775	16	5	88
3.	92	382	5		15
4.	39	206	3		6
5.	4	117			3
6.	14	93	1	2	10
7.		49			
8.		28	1		
9.		20			
10.		18			
11.		4			
12.	6	15			
13.		2			
14.		4			
16.		4			
18.		2			
19.		1			
21.		1			
22.		1			
24.		1			
28.		1			
29.		1			
36.		1			

*This is a sample of 12,897 people enrolled in Pennsylvania Blue Shield in 1980. Of these, 5,912 had no dental procedures in this period. These were removed from this table to match the WEA data more closely. (WEA data included only people who used one or more procedures in two years). Thus, the distributions represent the utilization of 6,985 people who had at least one dental procedure in one year. Only 3,772 of the patients had coverage for full-mouth scalings.

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Medical and Psychosocial Factors Predictive of Psychotropic Drug Use in Elderly Patients

L. Douglas Ried, PhD, Dale B. Christensen, PhD, and Andy Stergachis, PhD

Abstract: The purpose of this study was to investigate medical and psychosocial factors that may be used to identify patients at risk of psychotropic drug use. Population-based surveys were completed by 278 elderly health maintenance organization (HMO) patients in August 1984. Physical and mental health status and social support were measured in the survey. Automated prescription records from the year prior to and the year after the survey were linked to data from the survey. Patients received 737 prescriptions for psychotropic drugs during the two-year period under study. Doxepin (20.2 percent), flurazepam (15.2 percent), and diazepam (14.8 percent) were dispensed most frequently. Nearly 30 percent of the patients received a prescription for at least one psychotropic drug during the two-year period, and 14 percent received at least one prescription during both years. Three significant predictors of subsequent psychotropic drug use were: prior use (odds ratio = 17.2, 95% CI = 6.25, 47.33), the number of physical impairments (OR = 1.73, 95% CI = 1.05, 2.84), and the respondent's rating on the Alameda Health Scale (OR = 1.65, 95% CI = 0.99, 2.75). Patients' self-reported mental health status and sociodemographic characteristics were not significant predictors of subsequent use. (*Am J Public Health* 1990; 80:1349–1353.)

Introduction

Patients who are 65 years or older are at risk of incurring drug-related problems, in part, because they use drugs more frequently and more intensively than younger patients.[1] They are also more likely to have one or more chronic diseases that require drug treatment, often more intensive drug treatment.[2] Psychotropic drugs are among the most frequently consumed drugs and are among the most likely to cause problems for older patients.[3] National estimates indicate that 15 percent of ambulatory males and 24 percent of females over age 65 used a psychotropic drug in 1977.[4]

Once elderly patients begin using a psychotropic drug, use may continue for an extended period of time. One study reported daily use for 12 months or more by 38 percent of antidepressant users and 15 percent of antianxiety users.[5] Another study reported that about 60 percent of elderly patients who obtained a prescription for a psychotropic drug in the preceding year also received one in the subsequent year, and the highest prevalence of extended use (i.e., over 10 years) was among patients 65 years and older.[6]

Psychotropic drug use has been repeatedly demonstrated to be associated with physical and central nervous system side effects, including reduced mental functioning,[7] addiction,[8] sleep apnea and sleep disturbances,[9] and hip fractures.[10] Correlates of psychotropic drug use include female gender,[5,11–13] psychological and physiological conditions,[14–16] and lack of a supportive social network.[17,18] No definitive and persistent associations have emerged among marital status, economic status, and class differences in the use of mood modifiers.[11]

Most studies investigating psychotropic drug use have been retrospective and have relied primarily on self-reported psychotropic drug use.[5,15,19–21] Few studies have determined patients' psychosocial characteristics during the prior year and then examined their subsequent psychotropic drug use patterns.[12,22] The purpose of the present study was to examine a group of elderly patients for predisposing medical, physical, social, and psychological factors that explain and predict their subsequent psychotropic drug use.

Methods

The Setting

The study was conducted at Group Health Cooperative (GHC) of Puget Sound, a staff model, non-profit health maintenance organization (HMO) serving approximately 340,000 members in western Washington State. A population-based survey of an age and sex-stratified random sample of 1,432 GHC adult enrollees ages 18 years and older was conducted during August 1984. Enrollees over age 65 were oversampled in order to increase the accuracy of the estimates for this age group. A description of the survey has been reported earlier and includes a comparison of the sample with the Seattle Standard Metropolitan Statistical Area (SMSA) population.[23] Completed surveys were obtained from 79 percent of the original sample. The response rate increased to 81 percent after persons were removed from the denominator who were either deceased or who were no longer enrolled in GHC at the time of the survey. The sample for the present study was a subset of 278 persons over 65 years of age who completed the survey.

In addition to their survey data, prescription records for these elderly patients were obtained from GHC's automated prescription files. Data have been collected on each prescription filled at GHC since 1977 and includes individual patient's medical history number, drug name, strength, dosage form, therapeutic category, legend status, date dispensed, and prescriber. Psychotropic drugs were an out-of-pocket expense during the study period, provided to GHC enrollees at cost. Previous studies have shown the pharmacy files are very complete and accurately reflect the actual drug use experience of GHC enrollees.[24,25]

Subsequent Psychotropic Drug Use

Subsequent psychotropic drug use was determined by an examination of the prescription records for the one-year period after the survey (September 1, 1984 to August 30, 1985). Patients were classified as users or non-users of psychotropic drugs. Users were defined as patients obtaining one or more prescriptions for an antidepressant, anxiolytic, or sedative-hypnotic drug the subsequent year. Drug prod-

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From the School of Pharmacy, University of Washington, Seattle (all authors) and the Center for Health Studies, Group Health Cooperative of Puget Sound (Stergachis). Address reprint requests to L. Douglas Ried, PhD, Assistant Professor of Pharmacy, School of Pharmacy SC-69, University of Washington, Seattle, WA 98195. Dr. Christensen and Dr. Stergachis are both Associate Professors of Pharmacy. This paper, submitted to the Journal October 26, 1989, was revised and accepted for publication March 13, 1990.

ucts were classified according to the American Medical Association Drug Evaluations.[26] A second level of use was measured by counting the number of psychotropic drug prescriptions each patient received during the year after the survey, a simple indicator of the frequency or intensity of use.

Predictor Variables

Predictor variables included the patients' psychotropic drug use during the one-year period prior to the survey, sociodemographic characteristics, social support, and medical, psychological and physical health status. Sociodemographic characteristics included the patients' marital status, living situation, age, educational attainment, income, and gender. Mental health status and social support measures were obtained primarily from the Rand Health Insurance Survey[27] and included measures of depressive symptomatology, positive outlook, and closeness/completeness of relationships.

The first physical health status measure was based on whether patients reported having suffered from each of 17 specific medical conditions or physical ailments in the preceding 12 months (1 = yes; 0 = no); the measure was the sum of all the yes responses. Patients were also asked if their physical condition sometimes interfered with (1 = yes; 0 = no): a) dressing, feeding or moving around, b) climbing stairs or getting outdoors, c) ability to work because of some illness or injury, d) changing or cutting down the kind of work, or e) cutting down or causing them to stop any other activity such as hobbies or social activities; the total number of yes responses constituted this measure. Finally, self-reported health status was measured by asking patients to compare their own health to others' their age (1 = excellent; 2 = good; 3 = fair; and 4 = poor). Self-reported health status has been shown to be related to objective measures of health status.[28–30]

The Alameda Health Status Scale[31] was also used as a composite measure of physical health status. Its categories are based on questions related to disability, impairments, chronic conditions, symptoms and energy levels, i.e. severe (1) to no disabilities/impairments and a high energy level (7). Lower scores have been shown to be associated with poorer health outcomes (e.g., higher mortality risk and long-term psychotropic drug use).[31,32]

Statistical Analysis

Psychotropic drug users and non-users were compared on each of the predictor variables to determine if there were significant bivariate differences using *chi*-square and t-tests. Logistic regression techniques were used to derive an estimate of the natural logarithm of the odds of using psychotropic drugs in the next year because the predictor variables included both categorical and continuous variables[33–35]. Proportional effects were calculated from the logistic coefficients using the mean of the dependent variable as the point of comparison. Proportional effects are interpreted as the predicted change from the point of comparison in the proportion of cases belonging to category 1 (e.g. psychotropic drug users) rather than in category 0 (e.g. non-users) for a one-unit change in the predictor variable.[36,37] A residuals regression approach was taken. Predictor variables were entered into the equation in a hierarchical fashion in order to determine change in the likelihood of psychotropic drug use over time. Prior psychotropic drug use was entered first. The sociodemographic variables were entered in the second step and the psychosocial variables were entered in the final step. The improvement of fit in the logistic models was determined by evaluating the change in the log-likelihood statistic for each step. Finally, multiple regression was used to examine the relationship between the predictor variables and the number of psychotropic prescriptions obtained during the subsequent year.

Results

The 278 patients received 737 prescriptions for psychotropic drugs during the two-year period. Anxiolytic prescriptions accounted for 40.9 percent of the prescriptions dispensed (diazepam, 15.2 percent; meprobamate, 6.9 percent; alprazolam, 4.3 percent; all others, 14.5 percent). Antidepressant prescriptions were the next most frequently dispensed (35.8 percent) and included doxepin (20.8 percent), amitryptiline (9.8 percent), trazadone (3.5 percent) and all others (1.7 percent). Finally, sedative-hypnotics accounted for 23.3 percent of the psychotropic drug prescriptions dispensed (flurazepam, 15.6 percent; triazolam, 7.2 percent; pentobarbital, 0.5 percent).

Over 30 percent of the patients obtained at least one prescription for a psychotropic drug during the two-year period studied (Table 1). Nearly 23 percent of the patients had between one and 10 prescriptions filled, whereas over 8 percent of them received 11 or more prescriptions (range 0 to 39 prescriptions per patient). When only patients obtaining one or more psychotropic prescriptions were examined, they averaged nearly nine prescriptions per patient and over 60 percent of the patients that obtained a psychotropic prescription during the first year obtained another during the subsequent year. Nearly 53 percent of the patients who obtained a single prescription during the two-year study received an anxiolytic and 33 percent received a sedative-hypnotic. On the other end of the continuum, one patient received 39 psychotropic prescriptions (34 antidepressant prescriptions

TABLE 1—Frequency and Intensity of Psychotropic Drug Use

Prescriptions/Patient			Quantity/Prescription		
Prescriptions	No.	Percent	Pills	No.	Percent
None	193	69.4	10 or less	52	7.1
1–5	50	18.0	11–19	39	5.3
6–10	12	4.3	20–29	103	14.0
11–20	10	3.6	30–49	188	25.6
21–30	10	3.6	50–99	228	30.9
30 or more	3	1.2	100–199	90	12.2
			200 or more	37	5.0
TOTAL	278	100.0	TOTAL	737	100.0

and five anxiolytic prescriptions) and another patient received 36 prescriptions (all for anxiolytics).

Table 2 shows the results of the comparisons on each of the sociodemographic and health status predictor variables during the subsequent year between users and non-users of psychotropic drugs. The average user was nearly two years younger than the average nonuser. Psychotropic users reported that their physical condition interfered with more of their daily activities, had more self-reported medical conditions, had a less positive outlook on life, and, according to their own reports, were in somewhat poorer health. It is interesting to note that while self-reported health status of users was significantly lower than that of non-users, it was still in the "good" to "fair" range.

Multiple logistic regression results showed that prior drug use was the most important predictor of subsequent use (Table 3). A patient using a psychotropic drug during the year prior to the survey was 17.2 times (95% CI = 6.25 to 47.33) more likely to be a user during the year following the survey when compared to a non-user in the year before the survey. Stated another way, psychotropic drug use during the year before the survey increased the probability of being a user in the year after the survey to over 80 percent. The sociodemographic variables did not contribute to the prediction of subsequent use with the prior drug use variable in the model. For each additional daily activity interfered with, or for each additional unit increase on the Alameda Health Index, the probability of using a psychotropic drug increased 9.9 and 8.9 percent, respectively. The two mental health status variables (depressive symptomatology and positive outlook) were not significant when use in the prior year and physical health status were simultaneously controlled. Addition of the psychosocial and physical health status variables as a group did not significantly improve the fit of the equation over just prior use alone. In a separate analysis, the sociodemographic, psychosocial and physical health status measures were entered first into the equation (as a proxy for psychosocial and physical health status *prior* to psychotropic drug use in the year before the survey). Adding these variables first into the equation did not improve the fit of the model over the constant alone (change in *chi*-square goodness-of-fit = 5.17, 11 df). When drug use in the prior period was next added to the equation, the fit of the model improved (change in *chi*-square goodness-of-fit = 35.54, 1 df).

The only significant predictor of the number of psychotropic prescriptions obtained during the year following the survey was the number of psychotropic prescriptions obtained in the year before the survey (Table 4). The large positive relationship between the prior year's use and the subsequent year's use confirms that patients obtaining a larger number of prescriptions in the prior year are more likely to obtain a relatively larger number of prescriptions in the subsequent year.

Discussion

Our results show that elderly patients at the most risk of receiving a prescription for an anxiolytic, antidepressant or sedative-hypnotic drug in the following year were those who had used one in the previous year. Patients at greater risk were also more likely to be in somewhat poorer physical health and to have their daily activities interfered with to a somewhat greater extent. Patients' sociodemographic characteristics and measures of their mental health status were not significant predictors of subsequent psychotropic use. The frequency or intensity of subsequent drug use was also closely associated with the previous year's experience.

TABLE 2—Comparisons of Users and Non-users of Psychotropic Drugs in the Year after the Survey

Variables	Users (N = 59)	Non-users (N = 219)	Test Statistic	Difference	95% CI
Age	72.9	74.9	2.17[a]	1.98	0.09, 3.87
Percent female	66.1	54.8	1.98[b]	0.16	
Percent married	59.3	64.8	0.39[b]	0.52	
Percent living alone	27.1	28.3	0.00[b]	0.99	
Education (percent)			2.83[c]		
<8th grade	15.8	11.7			
9–11th grade	24.6	20.4			
High school graduate	24.6	26.3			
13–15 years	21.1	23.0			
College graduate	8.8	7.5			
Graduate school	5.3	11.3			
Income (percent)			3.12[c]		
<$10,000	26.5	23.5			
$10–$15,000	32.7	25.7			
$15–$25,000	22.4	29.9			
$25–$35,000	14.3	11.8			
$35–$50,000	2.0	5.9			
$50,000+	2.0	3.2			
Number of daily activities interfered with	1.04	0.61	2.09[a]	0.43	−0.03, 0.89
Number of medical conditions	4.32	3.42	2.13[a]	0.90	0.03, 1.79
Alameda scale	3.01	3.39	1.55[a]	0.38	−0.13, 0.88
Emotional ties	9.57	9.69	0.32[a]	0.12	−0.65, 0.99
Positive effect	43.89	46.77	1.91[a]	2.88	−0.21, 5.99
Depressive symptomatology	7.18	6.76	0.85[a]	0.42	−0.41,1.25
Self-reported health status	2.23	1.95	2.56[a]	0.28	0.02, 0.52

[a] Independent t-test
[b] *chi*-square (df = 2)
[c] *chi*-square (df = 5)

TABLE 3—Effects of Prior Use, Sociodemographic, and Psychosocial Characteristics on Subsequent Psychotropic Drug Use

Independent Variables	Step 1 β	Step 2 β	Step 3 β	% Predicted Proportional Change (Step 3)	Odds Ratio	95% Confidence Interval
Prior use	2.68*	2.66*	2.85*	61.1	17.2	6.25, 47.33
Sociodemographics						
Age	—	.033	.035	0.1	1.04	0.94, 1.14
Sex (female = 1)	—	.098	.008	0.2	1.01	0.31, 3.33
Married	—	.089	.093	1.5	1.10	0.14, 8.87
Living alone	—	.074	.100	1.6	1.11	0.14, 8.67
Physical and Psychosocial Characteristics						
Self-reported health status	—	—	.014	0.2	1.01	0.43, 2.41
Depressive symptomatology	—	—	.084	1.4	1.09	0.89, 1.33
Positive outlook	—	—	.008	0.1	1.01	0.92, 1.10
Emotional ties	—	—	.025	0.4	1.03	0.79, 1.33
Alameda scale	—	—	.501*	8.9	1.65	0.99, 2.75
No. of impairments	—	—	.549*	9.9	1.73	1.05, 2.84
No. of conditions	—	—	.087	1.4	1.09	0.84 ,1.42
Mean, Dependent	.1905	.1905	.1905			
Intercept	2.32	4.64	8.71			
chi-square goodness of fit (intercept only = 163.03)	126.87	126.18	119.32			
chi-square compared to intercept only:	36.16/1 df	36.85/5 df	43.71/12 df			
Change:	36.16/1 df	0.59/4 df	6.96/7 df			
p-value of change	p < .005	p = n.s.	p = n.s.			

*p < .05 coefficient/s.e. coeff which approximates t-distribution.

TABLE 4—Stepwise Multiple Regression of Predictor Variables on the Number of Prescriptions Obtained during the Subsequent Year

	Step 1 Beta	Step 2 Beta	Step 3* Beta	Standard Error B
Prior Use	0.67	0.67	0.64	0.05
Sociodemographics				
Age	—	0.08	0.07	0.02
Sex	—	0.01	0.01	0.22
Married	—	0.04	0.06	0.37
Living alone	—	0.06	0.06	0.38
Physical and Psychosocial Characteristics				
Self-reported health status	—	—	0.06	0.16
Depressive symptomatology	—	—	0.01	0.04
Positive outlook	—	—	−0.06	0.02
Emotional ties	—	—	0.03	0.06
Alameda scale	—	—	0.02	0.09
No. of impairments	—	—	0.05	0.11
No. of conditions	—	—	0.01	0.05
Adjusted R^2	0.45	0.45	0.45	
N	278	278	278	
F-ratio	223.0	45.4	19.3	

*Residuals analysis identified three outliers. The adjusted R^2 increased to 0.57 when the three outliers were eliminated.

Our finding that psychotropic drug use is independent of patients' physical and mental health status is consistent with the work of others.[38] There was also evidence of an association between physical disorders and the likelihood of any use, but not with the number of psychotropic prescriptions obtained during the subsequent year.

One reason for the predominant role of prior drug use as a predictor of later use, especially when compared to the findings of previous studies, may relate to the differences in data collection methods. In most of the previously cited retrospective studies, patients were asked to recall their psychotropic drug use during a period prior to collection of the psychosocial information. In terms of cause and effect, this more closely represents the notion that the psychotropic drug caused the psychosocial condition, not vice versa. In this study, we collected psychotropic drug use from prescription records before and after assessment of patients' sociodemographic and psychosocial characteristics.

Limitations to the study design exist that may constrain interpretation of the data. Psychotropic drug users were identified exclusively by automated pharmacy records. The GHC pharmacy data base provides a reasonably complete record of receipt of drugs; however, patients' actual consumption was unconfirmed. In addition, patients' sociodemographic characteristics and physical and mental health status data were obtained exclusively by self-reports.

Further work is needed to examine predictors of individual psychotropic drug categories. Small numbers of users prohibited such an analysis in this study. The sample size also influenced the time interval chosen for the present study. It would be beneficial to examine whether patient characteristics and mental health status were more important if the time interval were shorter than one year.

Another limitation of these data is the nature of the sample. The statistical findings are governed not only by its size, but also by the distribution of the variables within the sample. The distribution may differ in another sample. Therefore, caution should be used when generalizing our findings beyond the population of elderly patients from where our random sample was drawn.

Psychotropic drugs are often legitimately used for short-term treatment of physical disorders associated with normal aging; however, long-term use has been questioned.[39] Long-term psychotropic drug taking may be a learned behavior directed at helping elderly patients deal with chronic physical and emotional problems or reducing anxiety attacks and physical symptoms associated with withdrawal syndrome.[39–41] Prescription for a psychotropic drug may also serve to diminish physicians' frustration in treating the many ambiguous conditions associated with normal aging and to appear supportive to the patient.[38,42–43] Examination of the specific reasons for prescribing psychotropic drugs was beyond the scope of this study. However, we did document a pattern of continued use among a small portion of the elderly population that could not be explained or predicted by the level of patients' sociodemographic, mental and physical health status, and social support.

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The Association between Type A Behavior and Change in Coronary Risk Factors among Young Adults

Thomas F. Garrity, PhD, J. Morley Kotchen, MD, Harlley E. McKean, PhD, Diana Gurley, PhD,
and Molly McFadden, MS

Abstract: The association of Type A/B behavior pattern and changes in blood pressure, total serum cholesterol, serum triglyceride, body mass, and smoking was estimated in a cohort of 375 young Black and White men and women from a rural county in Central Kentucky between 1978–79 and 1985–88. Type A participants experienced significant increases in systolic (2.90 ± 1.29 mmHg) and diastolic (3.80 ± 1.17 mmHg) blood pressure and in cigarette smoking (3.26 ± 0.89 cigarettes per day) over the eight-year follow-up period, but Type B participants experienced no change. Type A and B individuals showed similar changes in total serum cholesterol, serum triglyceride, or body mass. Differences between behavioral types in blood pressure were present for women but not men, and for Blacks but not for Whites. These findings suggest a possible significance of the Type A pattern for the development of cardiovascular risk of young adults. (*Am J Public Health* 1990; 80:1354–1357.)

Introduction

Cardiovascular epidemiology has established blood pressure, serum cholesterol, and smoking as the most powerful and consistent predictors of coronary risk in middle age and older adults. Studies of children suggest that coronary risk factors are already present at early ages,[1,2] but relatively little epidemiologic data exist on the coronary risk status of young adults, especially those who have recently completed physical growth as they begin to enter typical adult roles in this society,[3] a phase characterized by changes in lifestyle and responsibility. The student becomes a worker, the single person a spouse and parent, with changes in eating, drinking, sleeping, and exercise.

The style of coping with these challenges may have a bearing on level of life stress and, through it, on markers of coronary risk. Type A, coronary-prone behavior pattern is believed to relate to risk for coronary heart disease.[4] We followed a sample of men and women from small town and rural backgrounds from an average age of 21 years to an average age of 29 to determine whether Type A/B behavior pattern is associated with change in coronary risk factors. Previously, we have shown clustering of risk factors in individuals in this population.[5]

Methods

The sample for the present study was derived from a 15-year longitudinal cohort study of young adults in a rural county of the Central Kentucky Bluegrass. In 1973, all high school students of Bourbon County, Kentucky (N = 1,142) were assessed for their blood pressures, body weights, and heights, as well as a few other demographic characteristics. Students' ages ranged from 14 to 19 years.

In 1978–79, a sample (N = 472) was drawn from the original study population to include the 20 percent of young adults who had the highest and lowest systolic blood pressure readings when they were teenagers in 1973. Because some of the subjects had not yet achieved adult growth at baseline, all of the subjects who were less than 16 years old in 1973 were also contacted for re-examination in 1978. Of the remaining group, 15 percent were selected at random for follow-up. Further description of this sample has been published previously.[5–7]

In 1985–88, the attempt was made to contact the entire population studied at baseline in 1973. Of the original population, 253 were lost to the study; 219 could not be located; 19 refused; 15 had died. Follow-up data were successfully gathered on 891 subjects, or 78 percent of the original pool.

The sample for the present study (N = 375) includes subjects for whom data are available both from the 1978 study and from the 1985–88 population follow-up. Subjects with very high or very low systolic blood pressure measurements when they were teenagers are overrepresented in this group, as are younger participants. As seen in Table 1, the typical 1985–88 participant was a White, married, full-time employed, high school graduate born in 1958. About half of the study group was male.

At both 1978–79 and 1985–88 data collections, participants were located by searches of local public documents, contacts with relatives and friends of the participants, and by placing classified advertisements in local and regional newspapers. Measurements were usually made at the project office in Bourbon County. For participants not willing to travel to the project office, home visits were performed by project staff throughout the state of Kentucky. For two out-of-state participants, Physical Measurements, Inc was hired to provide home visits at which a blood sample, blood pressure, and other anthropometric measurements were obtained according to protocol. Questionnaires were also completed by these two participants. Blood pressure, anthropometric, and medical history data were collected at both waves. The Type A measure (the Jenkins Activity Survey, Form C)[8] was administered only in 1985–88. Type A behavior was measured with the JAS Type A subscale comprising 21 items from the longer instrument. Each response was assigned a weight and summed according to the method of Jenkins.[8] Blood lipid analyses and other cardiovascular assessments[7] are available on a subset of the participants (N = 98).

Quetelet index[9]: Q = weight in kilograms/(height in meters)^2 was calculated. Blood pressure was measured in the right arm after the subject was seated for at least five minutes.

Address reprint requests to Thomas F. Garrity, PhD, Professor and Chairman, Department of Behavioral Science, College of Medicine Office Building, University of Kentucky, Lexington, KY 40536-0086. Dr. Gurley is also with that department at the University; Dr. Kotchen is with the Department of Medicine, West Virginia Medical Center; Dr. McKean is with the Department of Statistics, University of Kentucky; Ms. McFadden of Cleveland Heights, Ohio was with the West Virginia Medical Center at the time of the study. This paper, submitted to the Journal September 5, 1989, was revised and accepted for publication April 2, 1990.

TABLE 1—Distributions of Selected Demographic Characteristics of Participants, Bourbon County, Kentucky, 1978–79 and 1985–88 (N = 375)*

	1978-79		1985–88	
Age (years)	20.7 ± 1.8		28.6 ± 1.4	
Sex	N	(%)		
Male	200	(53.3)		
Female	175	(46.7)		
Marital Status			N	(%)
Single	251	(67.1)	101	(26.9)
Married	116	(31.0)	227	(60.5)
Separated	4	(1.1)	8	(2.1)
Divorced	3	(0.8)	39	(10.4)
Widowed	0	(0.0)	0	(0.0)
Employment				
Fulltime	90	(24.0)	283	(75.5)
Parttime	227	(60.5)	25	(6.7)
Not employed	58	(15.5)	67	(17.9)
Education				
Some High School	42	(11.2)	42	(11.2)
High School Diploma	243	(64.8)	232	(61.9)
Some College	83	(22.1)	53	(14.1)
College Diploma	7	(1.9)	48	(12.8)
Race				
White	301	(80.3)		
Black	74	(19.7)		

*375 is the total number of participants. Smaller numbers are due to missing information on some variables.

Three measurements by a trained observer were taken over a 10–15 minute period using a standard mercury manometer and cuff calibrated for arm circumference; these readings were averaged. Blood samples were analyzed for total serum cholesterol and serum triglyceride values by the autoanalyzer method described in the Lipid Research Clinics Manual of Laboratory Operations.[10]

We studied systolic and diastolic blood pressure, serum cholesterol, and serum triglyceride levels, body mass, and smoking behavior as indicators of risk. Change in risk status was measured as the difference in scores from 1978–79 to 1985–88 on each of these variables. Using change scores for each of the risk factors as outcomes, we calculated a three-way factorial model for all risk factors except smoking change, with behavior pattern, race and sex as the explanatory variables. Change in number of cigarettes smoked was treated as a categorical variable for log-linear modeling and contingency table analysis. Race, sex and body mass were included as potential control variables.[12] Relations between behavior pattern and risk variables were affected by race and sex, but not by change in body mass (data available on request) so change in body mass was removed from further analyses.

Results

Coronary risk factor means and standard deviations at the two time periods are shown in Table 2.

The three-way factorial analysis for change in systolic blood pressure revealed that only the main effect of behavior pattern accounted for significant variance in change in systolic blood pressure between 1978–79 and 1985–88 assessments. As shown in Table 3, Type A subjects experienced a mean 2.90 ± 1.29 (standard error) mmHg change, whereas Type B subjects showed a mean −1.47 ± 1.22 mmHg change in systolic blood pressure. The Type A subjects experienced greater increases in systolic blood pressure than Type B subjects in the entire sample ($p = 0.01$), in women ($p = 0.02$) but not in men ($p = 0.23$), and in Blacks ($p = 0.05$) but not in Whites ($p = 0.12$).

TABLE 2—Coronary Risk Factors in Participants, Means and Standard Deviations, Bourbon County, Kentucky, 1978–79 and 1985–88

	N	1978–79	1985–88
		Means SD	Means SD
Systolic Blood Pressure	375[+]	120.27 ± 14.90	120.69 ± 16.07
Diastolic Blood Pressure	375	73.08 ± 11.35	75.28 ± 11.88
Serum Cholesterol	98	167.21 ± 37.32	204.19 ± 43.18
Serum Triglyceride	98	107.53 ± 77.81	168.36 ± 99.65
Body Mass Index	371	24.17 ± 6.48	26.62 ± 5.81
JAS Type A Score	375		211.28 ± 72.82
Cigarettes/Day	369	7.24 ± 10.99	9.18 ± 13.39

[+] Fewer than 375 cases on a variable are due to missing information (except for the serum cholesterol and serum triglyceride variables in which sample design called for fewer cases).

In the three-way factorial analysis for change in diastolic blood pressure, none of the explanatory variables or their interactions were statistically significant, but Type A individuals rose an average 3.80 ± 1.17 mmHg, whereas Type B individuals rose 0.50 ± 1.11 mmHg in diastolic blood pressure (Table 4). The Type A subjects experienced somewhat larger increase in diastolic blood pressure than Type B subjects in each contrast especially among Blacks and women.

Factorial analysis of changes in total serum cholesterol, serum triglyceride, and body mass values showed no statistically significant main or interaction effects on change in any of these three risk variables.

Log-linear modeling analysis was performed with smoking, the outcome, divided into participants who decreased, remained unchanged and increased cigarette consumption from 1978–79 to 1985–88, and with behavior pattern, sex and race the predictors. Only behavior pattern was significantly associated with smoking change. Table 5 shows that Type As increased smoking more than type Bs, and that Type Bs held their smoking at the same level to a greater extent than Type As. Log-linear analysis found that the behavior pattern/smoking change association was statistically significant in Black females but not in the other three race/sex groups.

Discussion

In 1981, an authoritative panel examined the available evidence and judged the Type A behavior pattern to be a risk factor for coronary disease, comparable in impact to smoking, elevated cholesterol, and elevated blood pressure.[4] Subsequent results call that conclusion into question.[13–15] Recent research suggests that underlying hostility is the aspect of Type A behavior that is most strongly predictive of coronary heart disease endpoints.[16] A reanalysis of the Multiple Risk Factor Intervention Trial (MRFIT) data, one of the major recent studies failing to find a Type A/coronary disease association, has found indicators of hostility to significantly predict coronary disease incidence.[17]

Appropriate measurement of Type A behavior is uncertain,[18] but at best is only modestly associated with other coronary risk factors,[19,20] and has been seen as an independent risk factor. This study suggests that, in young adults, Type A behavior pattern is associated with more rapid increases in two well-accepted coronary risk factors. Type A behavior may operate differently at different ages to produce

TABLE 3—Mean Change (± standard error) in Systolic Blood Pressure for Sex/Race/Behavior-Pattern Groups*

	Males		Females		Both Sexes		
	Type A	Type B	Type A	Type B	Type A	Type B	A vs. B (p)
White	1.17 (1.38)	−0.66 (1.65)	1.72 (1.68)	−1.40 (1.58)	1.44 (1.09)	−1.03 (1.14)	0.12
Black	4.93 (3.49)	0.73 (2.88)	3.79 (3.10)	−4.56 (3.19)	4.36 (2.34)	−1.91 (2.15)	0.05
Both Races	3.05 (1.88)	0.04 (1.66)	2.76 (1.76)	−2.98 (1.78)	2.90 (1.29)	−1.47 (1.22)	
A vs B (p-value)	0.23		0.02				0.01

*Three-way analysis of variance found a significant ($p < .05$) main effect for Type A/B behavior pattern but no other main or interaction effects.

TABLE 4—Mean Change (± standard error) in Diastolic Blood Pressure for Sex/Race/Behavior-Pattern Groups*

	Males		Females		Both Sexes		
	Type A	Type B	Type A	Type B	Type A	Type B	A vs. B (p)
White	2.16 (1.26)	1.76 (1.50)	3.08 (1.53)	2.27 (1.44)	2.62 (0.99)	2.02 (1.04)	0.68
Black	4.40 (3.18)	0.00 (2.62)	5.58 (2.82)	−2.00 (2.90)	4.99 (2.13)	−1.00 (1.96)	0.04
Both Races	3.28 (1.71)	0.88 (1.51)	4.33 (1.60)	0.14 (1.62)	3.80 (1.17)	0.50 (1.11)	
A vs B (p-value)	0.29		0.07				0.04

*Three-way analysis of variance found no significant main or interaction effects.

TABLE 5—Contingency Table Analysis (chi-square) of the Association between Behavior Pattern and Change in Smoking Behavior, 1978–79 to 1985–88 (N = 369)

	Behavior Pattern			
	Type B		Type A	
Smoking Change	No.	%	No.	%
Decreased	21	11.8	28	14.7
Stayed Same	121	68.0	107	56.0
Increased	36	20.2	56	29.3
Total	178	100.0	191	100.0

chi-square = 5.76; d.f. = 2; p = 0.056

eventual coronary disease endpoints. In young adulthood, Type A might promote exacerbation of two coronary disease risk factors and thereby support early atherogenesis. In these younger adults, the Type A pattern may reinforce the smoking habit by motivating individuals toward greater cigarette consumption. The Type A pattern may also stimulate more rapid escalation of blood pressure in ways analogous to those demonstrated in stress laboratory research on Type A responsiveness to hostile challenges.[18] Type A behavior may become less strongly associated with these risk factors in middle age as life situations become more predictable and manageable.

Whether a Type A measure taken in 1985–88 reflects status in 1978–9 might be questioned, but Type A measures have been shown to be stable in young adults and in children.[21–23] At the least, these data suggest that those who measure as Type A in their late 20s are more likely to have experienced a steeper increase in blood pressure and smoking during the preceding decade than their Type B counterparts.

In the study of "Framingham offspring,"[3] their lifestyle, and changes in serum cholesterol, lipoproteins and blood pressure, no association was found between Framingham Type A measures (total scale, trait subscale, and job-related subscale) and changes in systolic or diastolic blood pressure, total, high density lipoprotein, low density lipoprotein, very low density lipoprotein cholesterols, and the ratio of total to high density lipoprotein cholesterols. However, in men only, the Framingham Type A trait subscale was independently predictive of the ratio of total to high density lipoprotein cholesterols by regression analysis. This significant finding could not be examined in the present study because of our lack of lipoprotein data. We are also unable to compare our body mass, serum triglyceride, and smoking results to those of the Framingham offspring study as the latter did not report Type A by body mass, serum triglyceride, and smoking associations.

Neither the design nor the data of the present study permits firm explanations for the stronger association of Type A with blood pressure and smoking changes in women and Blacks than in men and Whites. One possible explanation is that Type A women and Blacks, because of social discrimination, find themselves more consistently engaging in the Type A coping responses to challenges.

In summary, this nine-year cohort study uncovered evidence that individuals with the JAS-measured Type A behavior pattern were likely to experience a steeper increase in blood pressure and rate of cigarette smoking during their third decade of life than their Type B counterparts. These results were found in analysis of the whole study group, but appeared stronger in women and Blacks for blood pressure, and in Black women for smoking. In light of earlier findings of no relationships between Type A behavior and coronary disease risk factors, these results suggest that age differences may affect the strength of those associations.

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Behavioral Covariates of Waist-to-Hip Ratio in Rancho Bernardo

AMI LAWS, MD, RICHARD B. TERRY, BA, AND ELIZABETH BARRETT-CONNOR, MD

Abstract: We examined lifestyle and dietary habits in 685 men and 943 women (mean age 67 years) who completed an interview, examination, and food frequency questionnaire in 1984–87. Waist-to-hip ratio increased with age and body mass index in both men and women. In multiple regression, waist-to-hip ratio was independently associated with smoking, alcohol consumption, and exercise in men, and with smoking and alcohol consumption in women. The data suggest that waist-to-hip ratio is affected, at least in part, by behavioral, and potentially modifiable, factors. (*Am J Public Health* 1990; 80:1358–1362.)

Introduction

Male pattern abdominal or truncal obesity, as measured by either waist-to-hip girth ratio, or ratio of subscapular to triceps skin-fold, has been implicated in the development of ischemic heart disease,[1–3] stroke,[3,4] diabetes,[5,6] hypertension,[7–9] and all-cause mortality[2] in both men and women. The association is independent of overall adiposity. In fact, in at least one study, those at highest risk for subsequent ischemic heart disease and death were the leanest subjects, with the greatest waist-to-hip girth ratios.[2]

The etiology of abdominal obesity is not clear. Studies by Bouchard[10] have shown that heredity accounts for only 20–25 percent of the variance in waist-to-hip ratio in the general population. Shimokata[11] and we[12] recently reported that cigarette smoking is related to abdominal obesity independently of age and body mass index. However, since smoking is highly correlated with other health behaviors, such as diet, alcohol consumption, and physical activity, it was unclear whether the increased abdominal adiposity seen in smokers is due to smoking per se, or is a consequence of other, associated behaviors.

In this report we have attempted to determine whether other behaviors, specifically alcohol consumption, physical activity, and diet, also contribute to the waist-to-hip ratio in a defined population.

Methods

The population for this study consisted of older adult residents of Rancho Bernardo, a geographically defined community of Caucasian upper middle-class in Southern California. These subjects were first surveyed in 1972–74 as part of the Lipid Research Clinics Prevalence Study. As previously reported,[13] the initial response rate was 82 percent. In 1984–87, 81 percent of surviving men and women participated in a follow-up evaluation by telephone and/or clinic visit. The 76 percent who were seen in clinic form the basis of these analyses.

Smoking and alcohol use were determined by standardized, self-administered questionnaires which asked subjects if they had consumed any alcohol in the past year, the frequency of their consumption, and the specific forms of alcohol consumed in an average week and in the week prior to the survey. Grams of alcohol for the week prior to the survey and for an average week were computed by means of an algorithm.[14] For smoking, participants were asked if they had ever smoked cigarettes, whether they currently smoked, the usual number of cigarettes smoked per day, and number of years smoked. Exercise was assessed by asking participants if they regularly exercised three times per week. Diet was assessed in a subsample of 737 participants with the Harvard-Willett semi-quantitative food frequency questionnaire.[15]

In clinic, a nurse measured height and weight with subjects in light clothing, without shoes. Body mass index (weight in kg/height in m^2) was calculated as an estimate of obesity. Waist and hip girths were measured over light, single thickness clothing with participants standing in an erect position with feet together. Participants were asked to bend forward and return to an erect posture. Waist girth was measured in cms at the bending point. Hip girth was measured in cms at the iliac crest. A 75 gram oral glucose tolerance test was administered in the morning after a requested 12-hour fast; diabetes was defined by World Health Organization (WHO) criteria.[16]

Among the 841 men and 1,117 women ages 50–79 who participated, 125 men and 114 women had non-insulin dependent diabetes by history or glucose tolerance test. Non-insulin dependent diabetes, which is strongly genetically determined,[17] and related to waist-to-hip-ratio was excluded from the present analyses in order to emphasize attributes which could be modified by behavior. Subjects with incomplete diabetes data (N = 84) were likewise excluded. These analyses, therefore, are based on 685 men and 943 women without diabetes. There were no other exclusions for disease, but all subjects were healthy enough to come to the clinic.

Data were analyzed using Statistical Analysis Systems (Cary, NC). Groups were compared using analysis of variance; age and age and body mass index-adjusted values for waist-to-hip ratio were compared by analysis of covariance. Pearson correlation coefficients were used to assess the correlation of waist-to-hip ratio with the dietary variables. Waist-to-hip ratio was sufficiently normally distributed not to require transformation. Multiple linear regression was used to determine the independent contributions of the behavioral variables to waist-to-hip ratio. All probabilities reported are for two-tailed tests.

Results

The mean age of both men and women was 67 years (Table 1). Men were heavier and had a higher waist-to-hip ratio than women. BMI decreased with age in men ($r = -.19$, $p < .001$) but not in women. Waist-to-hip ratio correlated with BMI in both men and women, ($r = .40$; $r = .31$, respectively).

From the Stanford University School of Medicine (Laws is a post-doctoral fellow, Terry is a pre-doctoral fellow) and the Department of Community and Family Medicine, University of California/San Diego (Barrett-Connor). Address reprint requests to Elizabeth Barrett-Connor, MD, Professor and Chair, Department of Community and Family Medicine M-007, School of Medicine, University of California/San Diego, La Jolla, CA 92093. This paper, submitted to the Journal October 31, 1989, was revised and accepted for publication April 13, 1990.

TABLE 1—Clinical and Behavioral Characteristics of Men and Women Ages 50–79 in Rancho Bernardo

	Men N = 685	Women N = 943
Age (mean years)	67.5 (8.3)	67.6 (8.2)
BMI (mean kg/m^2)	26.2 (3.3)	24.4 (3.7)
Waist-to-Hip Ratio	0.915 (.050)	0.789 (.062)
Smokers		
% Never	28	47
% Former	60	37
% Current	12	16
Cigarettes per day* (mean)		
Former	26 (16)	16 (15)
Current	24 (17)	16 (13)
% Drinkers	75	65
Alcohol past week** (mean gms)	142 (126)	102 (92)
Alcohol average week** (mean gms)	152 (123)	111 (88)
% Exercisers	87	80

Values are mean ± SD.
*among smokers
**among drinkers

When adjusted for BMI, waist-to-hip ratio increased with age in both men and women (r = .11, r = .23, respectively).

Almost two-thirds of the men and over one-third of the women had smoked in the past, but only 12 percent of the men and 16 percent of the women were current smokers. Over two-thirds of men and women reported drinking at least one or two drinks per month and most reported regular physical activity. Among users, men smoked more cigarettes and drank more alcohol than women.

Waist-to-hip ratio was higher in men and women who had ever smoked compared to those who had never smoked, and was higher in the upper quartile of drinkers (>9 oz/wk for men; >7 oz/wk for women) compared to all others. Waist-to-hip ratio was lower in men and women who reported regular exercise compared to those who did not (Table 2).

Waist-to-hip ratio, adjusted for age and BMI, was lowest in never smokers, intermediate in former smokers, and highest in current heavy smokers (Figure 1). The magnitude of the difference in waist-to-hip ratio between never smokers and current heavy smokers was as large in the women as in the men, despite the fact that within each category the women smoked fewer cigarettes per day.

Figure 2 shows mean waist-to-hip ratios by five drinking categories based on alcohol consumption in the week prior to the clinic visit week. Similar results (not shown) were found based on alcohol consumption in an average week. In both men and women, the heaviest drinkers had the highest waist-to-hip ratios.

Correlations between a number of dietary variables from the Harvard-Willett food-frequency questionnaire and waist-to-hip ratio were examined in 335 men and 402 women, shown in Table 3. Among the dietary variables, percent calories from alcohol was positively correlated with waist-to-hip ratio in men and women, as was absolute alcohol intake (data not shown). The only other dietary variable significantly correlated with waist-to-hip ratio was percent calories from carbohydrate (exclusive of alcohol) which was inversely correlated with waist-to-hip ratio in men.

Cigarettes per day for former and current smokers, alcohol as a continuous variable, and exercise were entered into a multiple linear regression model with age and BMI. There was no significant interaction of smoking with alcohol intake in the model. As shown in Table 4, in both sexes, age, BMI, alcohol intake, and smoking were independently associated with waist-to-hip ratio. Exercise was independently associated with waist-to-hip ratio in men but not in women. Overall, these attributes explained 21 percent of the variance in waist-to-hip ratio in men and 16 percent in women.

Discussion

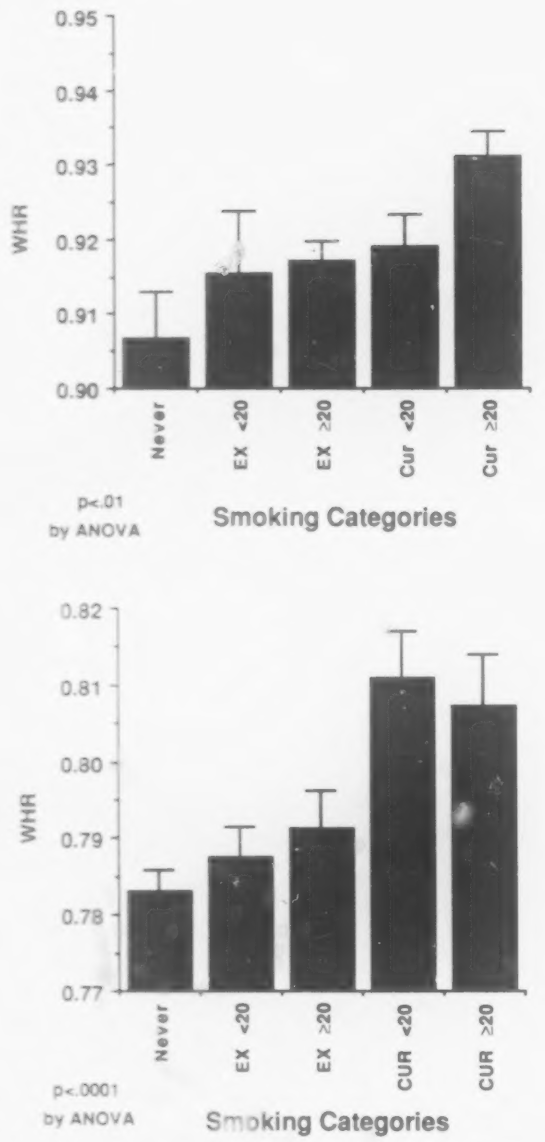
The study confirms observations by Shimokata,[11] and extends earlier observations in this population[12] by demonstrating that effects of smoking and drinking on waist-to-hip ratio are independent and additive.

The fact that the increase in waist-to-hip ratio associated with drinking was accounted for by the heaviest drinkers is consistent with one possible mechanism, an increase in adrenal cortisol production which has been reported in alcoholics and has been termed "alcohol-induced pseudo-Cushing's syndrome."[18,19] It is well-known that central obesity is a prominent feature of Cushing's syndrome.[20] Although pseudo-Cushing's is apparently a relatively uncommon complication of alcoholism,[21] it is possible that a lesser degree of abnormality in cortisol secretion associated with heavy drinking could affect waist-to-hip ratio in the absence of clinical Cushing's syndrome.

The association of smoking with increased waist-to-hip ratio in women may be explained by the fact that women smokers have increased levels of adrenal androgens[22,23]

TABLE 2—Mean Waist-to-Hip Ratio by Presence or Absence Behavioral Variables

	Men			Women		
	N = 685			N = 943		
Behavior	Presence	Absence	Difference (95% CI)	Presence	Absence	Difference (95% CI)
Smoker (ever)						
Unadjusted	.918	.906	.012 (.004, .020)	.794	.784	.010 (.002, .018)
Age adjusted	.918	.906	.012 (.004, .020)	.794	.784	.010 (.003, .018)
Age and BMI adjusted	.918	.906	.012 (.004, .019)	.795	.783	.012 (.004, .019)
Drinker (upper quartile vs all others)						
Unadjusted	.928	.912	.016 (.007, .026)	.804	.786	.018 (.007, .028)
Age adjusted	.928	.912	.016 (.007, .026)	.803	.787	.016 (.006, .026)
Age and BMI adjusted	.929	.912	.017 (.008, .026)	.803	.787	.016 (.007, .026)
Exerciser (regular)						
Unadjusted	.912	.933	−.021 (−.010, −.032)	.787	.797	−.010 (−.001, −.020)
Age adjusted	.912	.934	−.022 (−.010, −.033)	.787	.797	−.010 (−.001, −.020)
Age and BMI adjusted	.913	.930	−.017 (−.007, −.027)	.789	.797	−.004 (.005, −.013)

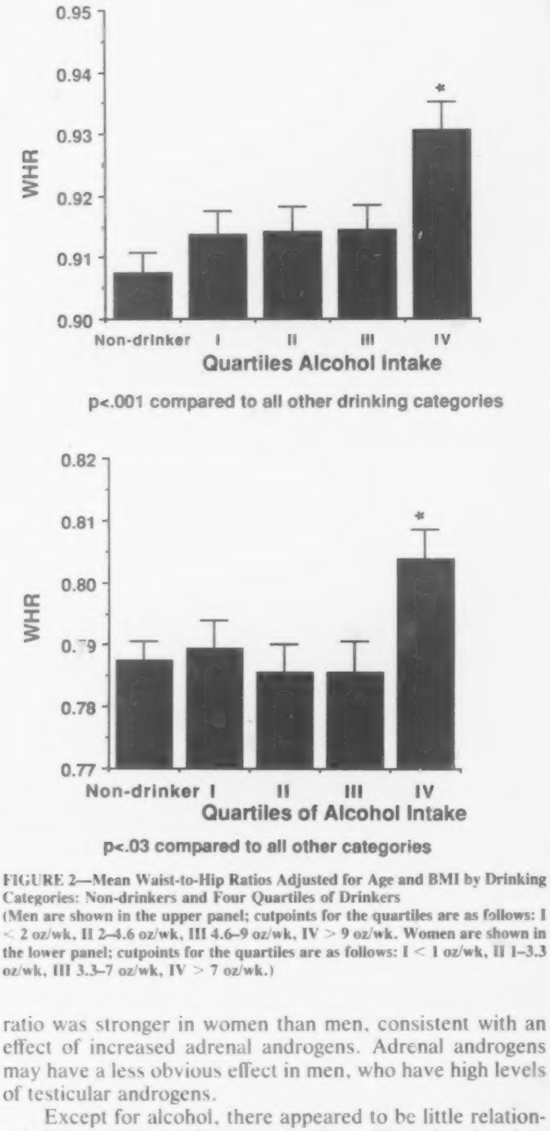


p<.01
by ANOVA

p<.0001
by ANOVA

FIGURE 1—Mean Waist-to-hip Ratios Adjusted for Age and BMI by Category of Smoking: Never Smokers, Exsmokers of <20 Cigarettes per Day, Exsmokers of ≥20 Cigarettes per Day, Current Smokers of <20 and ≥20 Cigarettes per Day (Men are shown in the upper panel; women in the lower.)

which have been shown to be associated with upper body obesity.[24] Increased cortisol levels in postmenopausal women smokers[23] also may contribute to increased abdominal obesity. In men, smoking is associated with increased adrenal androgens and estrogen, however, there are few data on the association of sex hormones with waist-to-hip ratio in men.[25] One study showed the only difference between men with higher vs lower waist-to-hip ratio was a lower sex hormone binding globulin in the former group.[26] In the data presented here, the association of smoking with waist-to-hip



p<.001 compared to all other drinking categories

p<.03 compared to all other categories

FIGURE 2—Mean Waist-to-Hip Ratios Adjusted for Age and BMI by Drinking Categories: Non-drinkers and Four Quartiles of Drinkers (Men are shown in the upper panel; cutpoints for the quartiles are as follows: I < 2 oz/wk, II 2–4.6 oz/wk, III 4.6–9 oz/wk, IV > 9 oz/wk. Women are shown in the lower panel; cutpoints for the quartiles are as follows: I < 1 oz/wk, II 1–3.3 oz/wk, III 3.3–7 oz/wk, IV > 7 oz/wk.)

ratio was stronger in women than men, consistent with an effect of increased adrenal androgens. Adrenal androgens may have a less obvious effect in men, who have high levels of testicular androgens.

Except for alcohol, there appeared to be little relationship between dietary factors and waist-to-hip ratio other than the inverse relationship with carbohydrates in men. This association could be a chance finding, reflecting the multiple diet comparisons. Alternatively, a high carbohydrate diet may be associated with a healthy lifestyle and exercise.

The apparent sex difference in the relationship of exercise to waist-to-hip ratio is compatible with metabolic studies which show different effects of exercise on fat mass in men compared to women.[27] After a physical training program, obese men show a decrease in total body fat, whereas women do not.[28–30] Since waist-to-hip ratio is highly correlated with percent body fat,[31] exercise could have a greater effect on waist-to-hip ratio in men than in women.

These findings are based on older, White men and women, and may not be generalizable to younger sub-

TABLE 3—Correlations of Waist-to-Hip Ratio with Dietary Variables as Measured by the Harvard-Willett Food Frequency Questionnaire

	Men N = 335	Women N = 402
Total calories		
Unadjusted	.04	.02
Age and BMI Adjusted	.10	.01
% Calories from Alcohol		
Unadjusted	.19§	.15#
Age and BMI Adjusted	.25**	.18†
% Calories from Carbohydrate		
Unadjusted	−.25**	−.06
Age and BMI Adjusted	−.25**	−.03
% Calories from Fat		
Unadjusted	.04	−.08
Age and BMI Adjusted	.01	−.11
% Calories from Protein		
Unadjusted	.08	−.02
Age and BMI Adjusted	.00	−.03
Fiber, dietary		
Unadjusted	−.14*	−.08
Age and BMI Adjusted	−.09	−.00
Caffeine		
Unadjusted	.09	−.05
Age and BMI Adjusted	.07	−.01
Sucrose		
Unadjusted	−.06	−.06
Age and BMI Adjusted	−.05	−.04

*p = .0114, #p = .0023, †p = .0018, §p = .0005, **p = .0001

TABLE 4—Multiple Linear Regression of Behavioral Variables on Waist-to-Hip Ratio

	Men*		Women**	
	β	Standard Error	β	Standard Error
Age (years)	.00066	0.0002	.00173	0.0002
BMI (kg/m^2)	.00618	0.0005	.00537	0.0005
Alcohol (mean gms/past week)	.00005	0.00001	.00006	0.00002
Cigarettes (mean no.)	.00027	0.0001	.00034	0.0001
Exercise (Y/N)	−.01520	0.0052	−.00288	0.0046

*R^2 = 0.21
**R^2 = 0.16

jects or to minorities. However, the age group studied is that in which "middle-aged spread" is classically manifest. The analyses were based on self-reported behaviors, and it is possible that these were inaccurately reported by the participants. Probable validity of the reported alcohol intake is strengthened by concordance with aspartate aminotransferase activity and HDL-cholesterol in this cohort.[32] Likewise, the validity of self-reported exercise in this cohort is strongly suggested by its concordance with pulse and HDL-cholesterol.[33] In any event, misclassification of self-reported behaviors usually serves to obscure an association, not create one.[34]

The fact that waist-to-hip ratio appears to be determined, at least in part, by behavioral factors has implications for intervention. This risk factor for increased morbidity and mortality may be modifiable by interventions aimed at changing the behaviors which affect it. This would be expected to have a favorable impact on general, and specifically, cardiovascular health.

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The Framingham Disability Study: Relationship of Various Coronary Heart Disease Manifestations to Disability in Older Persons Living in the Community

JOAN L. PINSKY, MA, ALAN M. JETTE, PT, PHD, LAURENCE G. BRANCH, PHD,
WILLIAM B. KANNEL, MD, MPH, AND MANNING FEINLEIB, MD, DRPH

Abstract: The relation between coronary heart disease and disability was examined in 2,576 community-dwelling women and men ages 55–88 years. These Framingham Study participants were originally recruited in 1948–51 for an examination of cardiovascular disease. Twenty-seven years later, remaining members of the cohort were interviewed to ascertain physical abilities, and a score on a disability scale was assigned. Multivariate logistic analyses examined disability in relation to uncomplicated angina pectoris (AP), complicated AP, and coronary heart disease other than AP, controlling for possible confounders. In younger and older women and men, uncomplicated and complicated AP were associated with disability. Coronary heart disease other than AP was associated with disability only in the younger men. Congestive heart failure predicted disability only in the women. These results suggest that onset of AP should be recognized as a critical point in the development of disability and that AP is a better predictor of disability than is myocardial infarction or coronary insufficiency. (*Am J Public Health* 1990; 80:1363–1368.)

Introduction

Although the coronary heart disease (CHD) mortality rate in the United States has been declining for two decades,[1] CHD remains the leading cause of death.[2] There are indications that CHD morbidity has not been following the same trend as mortality, and that the number of persons, particularly older persons, disabled from CHD remains high. Estimates of the percentage of persons with ischemic heart disease who were disabled by their condition in 1980 range from 36 percent in 55–64 year old women to 76 percent in men 75 years and older.[3] Heart disease ranks as third among the leading chronic conditions causing limitation of activity; in 1983–85, 5.6 million persons were affected* and this number was higher than a 1980 estimate.**

A better understanding of the CHD-disability relationship may provide leads to primary and secondary preventive health care strategies.[4] Some researchers have sought to identify predictors of disability among persons with CHD.[5–7] The Framingham Study enables us to look at disability in a non-institutionalized population of older persons, to compare physical limitations of those with and without CHD, and to learn more about the relationship of disability to specific CHD manifestations. Our specific question is: among a population of non-institutionalized women and men ages 55–88 years in 1976–78, is disability related to uncomplicated angina pectoris (AP), CHD other than AP, and AP complicated by other CHD? Since congestive heart failure is closely associated with CHD, we will also examine its relation to disability.

Methods

Study Sample

Between 1948 and 1951, the Framingham Study assembled a cohort of 5,209 persons, ages 28–62 years in Framingham, Massachusetts.[8] Information on potential cardiovascular disease risk factors and disease manifestations were obtained every two years by interview, physical examination, and laboratory tests. By December 1979, 31 years after the inception of the study, the 15th examination cycle had been completed.

The Framingham Disability Study (FDS) was designed to yield a quantitative assessment of levels of physical and social disability based on respondents' self-reported information. Surviving members of the Framingham Study reported their current level of functioning in interviews from September 1976 (midway through the 14th examination cycle) through November 1978 (midway through the 15th examination cycle). This time period will be referred to throughout this report as examination 14/15. Descriptions of the prevalence findings on physical disability from this study have been reported previously.[9]

When the Framingham Disability Study interviewing began, 1,649 of the original cohort of 5,209 had died, 218 had moved away and dropped out of the study, 151 had refused to return for examination, and 354 had been missing from the study for no known reason for at least five years.

Of the remaining 2,837 individuals, 183 (6 percent) were not included in the sample of FDS participants: 41 were nursing home residents who were unable to come to the clinic and were therefore excluded; 87 had come into the clinic for their 14th examination before the disability interviewing began and had not come in for their 15th examination by the time the disability interviewing ended; 43 refused to be interviewed; 11 came for their physical examination but were missed by the disability interviewer due to scheduling difficulties and were unable to be rescheduled; and one person died before a rescheduled interview was taken, leaving 2,654 remaining. Members of the original cohort who did not

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Address reprint requests to Joan L. Pinsky, MA, Division of Epidemiology and Clinical Applications, National Heart, Lung and Blood Institute, National Institutes of Health, Federal Building, Rm 2C08, Bethesda, MD 20892. Dr. Jette is with the Institute of Health Professions, Massachusetts General Hospital, Boston; Dr. Branch is with the Gerontology Center, Boston University; Dr. Kannel is with the Department of Medicine, Boston University School of Medicine, Section of Preventive Medicine and Epidemiology; Dr. Feinleib is Director of the National Center for Health Statistics. This paper, submitted to the Journal July 12, 1989, was revised and accepted for publication March 30, 1990.

*Unpublished table: C3-8501, Prevalence and rate per 1,000 of chronic conditions of six chronic condition lists and ICD codes: United States, 1980. National Health Interview Survey, Center for Health Statistics, November 7, 1986.

**Unpublished table: C-8003, Prevalence of published chronic conditions rates by sex and age, and percent and rates for selected health related characteristics: United States, 1980. National Health Interview Survey, National Center for Health Statistics, undated.

participate in the FDS for all of the above-mentioned reasons other than death (N = 882), when compared with FDS participants on characteristics at entry to the Framingham Study, were more likely to be female, older, have a higher mean systolic blood pressure, and a higher body mass index (Table 1).

Seventy-eight (3 percent) of the 2,654 participants were excluded from present analyses for the following reasons: 76 persons had missing values in one or more of the items used to calculate a score on the disability scale, and two had missing values in one of the risk factor measures used as covariates in the analyses. There remained 1,511 women and 1,065 men available for this investigation. Ages ranged from 55 to 88 years for the women and from 56 to 86 years for the men.

Definition of Variables

Coronary Heart Disease Manifestations and Congestive Heart Failure: Determination of these end points was based on information from biennial examinations, surveillance of hospital discharge records and information obtained from private physicians. A panel of physician investigators reviewed all pertinent information against minimal uniform criteria before final assignment of a diagnosis.[10] Examination 14/15 disease status was determined using diagnoses made at any time from examination 1 through examination 14/15.

Myocardial Infarction was diagnosed on the basis of unequivocal specific electrocardiographic criteria or elevated serum enzyme levels. This classification included both recognized myocardial infarctions (ones in which onset of symptoms led the individual to seek medical attention) and unrecognized myocardial infarctions (ones in which no symptoms occurred prior to an examination by a physician).

Angina Pectoris: Brief recurrent chest discomfort of up to 15 minutes duration precipitated by exertion or emotion and relieved by rest or nitroglycerin was regarded as angina pectoris (AP) if two physicians interviewing the participant agreed that AP was definitely present. Electrocardiographic abnormalities were not taken into account in making the diagnosis of AP.

Coronary Insufficiency was diagnosed when a history of ischemic chest pain lasting one hour or more was associated with transient ischemic electrocardiographic criteria.

Congestive Heart Failure: A definite diagnosis of congestive heart disease required a minimum of two major criteria (e.g. increased heart size and rales) or one major and two minor criteria (e.g., rales, bilateral ankle edema, and night cough).

TABLE 1—Selected Characteristics of Original Framingham Cohort Members by Participation in the Framingham Disability Study

	Participation Status		
Characteristics at Entry	No*	Yes	95% CI of difference
Number of persons	882	2,654	
Percent female	65.3	58.9	2.72, 10.03
Age (years)	44.0	40.9	2.49, 3.65
Systolic blood pressure (mmHg)	135.2	131.2	2.55, 5.50
Body mass index (kg/m^2)†	25.7	25.4	0.02, 0.64
Cigarettes per day‡	9.2	8.4	−0.02, 1.72

*Persons who died prior to the start of the Framingham Disability Study were excluded from this group.

Number of observations with missing values: † = 4; ‡ = 19.

The following mutually exclusive subgroups of CHD were considered: uncomplicated AP (AP unaccompanied by coronary insufficiency or myocardial infarction); CHD other than AP (coronary insufficiency or myocardial infarction); and complicated AP (AP accompanied by coronary insufficiency, myocardial infarction, or both).

Disability Scale: The measure of disability used was based on three questions drawn from the Rosow and Breslau Functional Health scale:[11] "Are you able to do heavy work around the house, like shoveling snow or washing windows, walls, or floors, without help?" "Are you able to walk up and down stairs to the second floor without help?" and "Are you able to walk half a mile without help? That's about 8 ordinary blocks." The scale ranged from 0 to 3, reflecting the number of activities which the respondent could not do without help. In other words, the higher the score, the more areas of disability were reported; a score of 0 indicated no disability.

Control Variables: To minimize the chance of detecting a spurious relation between the onset of CHD and subsequent disability, we controlled for the effects of co-existing conditions: other cardiovascular diseases, cardiovascular risk factors, and reported health history of other diseases related to disability. Presence of cardiovascular diseases other than CHD, namely stroke and intermittent claudication, were determined in a manner identical to that used to identify CHD and congestive heart failure, by a panel of physicians using uniform criteria after each biennial examination. Detailed descriptions of clinical criteria employed for the diagnoses of these diseases are available elsewhere.[10]

Cardiovascular Disease Risk Factors: We controlled for the long-term presence of hypertension (coded as a categorical variable taking into account anti-hypertensive drug use as well as blood pressure levels), body mass index (weight, kg, divided by height, m^2), cigarette smoking, and blood sugar, the four cardiovascular disease risk factors found in preliminary analysis to be related to the disability scale. For an explanation of how these risk factor measures were calculated, see Pinsky, *et al.*[12]

Other Health Problems: Several open-ended questions solicited information on participants' lifetime health status. Participants were asked first to mention all of the serious health problems they had had over the years, and were told to include "things that were serious at the time, things that remain as potential problems, or things that in any way limit activities at the present time." A two-part question followed: respondents were asked to mention things they used to do that the cannot do now or had to stop or cut down on; and they were asked what health problem caused each limitation. A maximum of 11 self-reported health problems for each individual were noted during the interview and subsequently classified into one of 233 codes. For the present analyses, 131 codes were grouped into 16 categories of health problems. Age-adjusted analyses indicated that six of the 16 categories were related significantly to the disability scale and are included as control variables in these analyses: amputations, arthritis, back problems, gastrointestinal problems, neurological problems (excluding stroke), and respiratory problems. The categories of health problems which were not related to the disability scale were allergies, cancer, ear problems (including hearing impairment), eye problems (including visual impairment), kidney problems, liver problems, hernias, bone injuries and other accidents, and bone problems other than accidents (data available on request to author).

Statistical Analysis

We chose a logistic model for multivariate analysis which is similar to that described by McCullagh, since the disability scale, our dependent variable, is an ordered categorical variable consisting of four levels of increasing severity. In this approach, the probability of being at the ith level or higher is written as a logistic function of the explanatory variables with a indicator term depending on the value of i. A modification of McCullagh's analysis permits inclusion of continuous explanatory variables. Maximum likelihood was used for estimation, and odds ratios were calculated as the exponent of the appropriate logistic regression coefficient. This model assumes that the underlying odds ratios for the three ordered comparisons of scores on the dependent variable (1 versus 0, 2 versus 1, and 3 versus 2) are equal. We used the logistic procedure in the Statistical Analysis System available from the SAS Institute.[14]

In a previous report, a different disability scale and disease were analyzed separately for women and men, controlling for age.[15] In this report, separate analyses were conducted for two age categories among women and men in an effort to provide additional insights. Furthermore, the lack of a linear relationship between disability and age[9] suggests the relevance of conducting age-specific analyses, rather than putting age in the multivariate model. Each regression included the following control variables: cardiovascular risk factors (i.e. hypertension, cigarette smoking, body mass index, and blood sugar); intermittent claudication and stroke; and six other health problems, (i.e. amputations, arthritis, back problems, gastrointestinal problems, neurological problems, and respiratory problems).

Results

Mean age was 68.3 years for 1,511 women and 67.2 years for 1,065 men. The percentage distribution of scores, from 0 to 3, on the disability scale were, respectively, 61.0, 26.6, 8.2, and 4.2 for women and 78.1, 16.6, 3.6, and 1.7 for men. Not only did a greater percentage of women report disabilities compared to men, but they reported being disabled in more activities than men.

Table 2 displays the age- and gender-specific prevalence of scores greater than 0 on the disability scale by CHD and congestive heart failure status. In all gender and age strata, disability was more prevalent in persons with CHD or congestive heart failure than in persons free of CHD or congestive heart failure. Persons with AP were more likely to be disabled than those with CHD manifestations other than AP. Congestive heart failure was more disabling in women than men.

Gender- and age-specific multivariate logistic analyses examined the relationship between the disability scale and the three CHD manifestations and congestive heart failure. Using the results of these analyses, odds ratios were calculated, representing a comparison of the disability status of persons with the manifestation to persons free of all disease categories included in the model (Table 3). An odds ratio of one means that those with the manifestation have the same chance of having a higher score on the disability scale as persons free of all disease categories included in the model, within each particular gender and age subgroup. For both women and men, AP both uncomplicated and complicated is strongly related to the disability scale in both age groups, although the relationship weakens with age. CHD other than AP is related to the disability scale only in the younger men. Congestive heart failure is related to disability only in the younger women.

Table 4 presents the results of analyses similar to those in Table 3, except the model also includes terms for other cardiovascular diseases, cardiovascular risk factors, and the self-reports of non-cardiovascular health conditions. The inclusion of these additional covariates did little to affect the results of the analyses. When comparing Table 4 to Table 3, one notes only minor shifts, most of which are downward, in the magnitude of the relationships between disability and disease.

Myocardial infarction is often the event of interest in epidemiological studies and is a much more frequent diagnosis when compared to coronary insufficiency. If we exclude from our pool of persons under study the 56 with coronary insufficiency, our CHD categories become uncomplicated AP, uncomplicated myocardial infarction, and AP complicated by myocardial infarction. Analyses based on these definitions resulted in findings similar to those which included coronary insufficiency. When the three levels of disability were combined, making the dependent variable dichotomous, the pattern of significant relationships did not change at all. For most associations, the strength of the relationship was greater when compared to the analyses using the multichotomous measure.

Discussion

Although our results indicate that AP is a better predictor of disability than other CHD, we were unable to control

TABLE 2—Percent of Women and men with Disability* at Examination 14/15 by Coronary Heart Disease and Congestive Heart Failure Status and Age

	Women			Men			Women			Men		
	N	% with Disability	95% CI†	N	% with Disability	95% CI	N	% with Disability	95% CI	N	% with Disability	95% CI
	Ages 55–69 years			Ages 56–69 years			Ages 70–88 years			Ages 70–86 years		
No Coronary Heart Disease or Congestive Heart Failure	829	24.7	21.8, 27.7	574	9.4	7.0, 11.8	471	48.8	44.3, 53.3	275	26.9	21.7, 32.2
Uncomplicated Angina Pectoris	44	65.9	51.9, 79.9	39	46.2	30.5, 61.8	64	81.3	71.7, 90.8	38	50.0	34.1, 65.9
Coronary Heart Disease other than Angina Pectoris	14	50.0	23.8, 76.2	40	32.5	18.0, 47.0	26	61.5	42.8, 80.2	33	39.4	22.7, 56.1
Complicated Angina Pectoris	23	69.6	50.8, 88.4	42	66.7	52.4, 80.9	19	94.7	84.2, 100	21	66.7	46.5, 86.8
Congestive Heart Failure	15	80.0	60.0, 93.3	7	42.9	14.2, 71.4	25	88.0	80.0, 100	14	57.1	31.2, 83.1

*Defined as a score of 1, 2, or 3 on the disability scale.

†CIs were calculated as follows: when both n × p and n × q were greater than 5, the normal approximation was used; otherwise the exact binomial was used.

TABLE 3—Association between the Disability Scale and Coronary Heart Disease Manifestations and Congestive Heart Failure by Sex and Age, from Logistic Regression Analyses (Terms in the model: the four disease categories listed below and age)

	Women		Men		Women		Men	
Disease Category	Odds Ratio	95% CI	Odds Ratio	95% CI	Odds Ratio	95% CI	Odds Ratio	95% CI
	Ages 55–69 years (N = 918)		Ages 56–69 years (N = 695)		Ages 70–88 years (N = 593)		Ages 70–86 years (N = 370)	
Uncomplicated Angina Pectoris	5.87	3.30, 10.43	6.71	3.43, 13.11	2.27	1.39, 3.70	2.56	1.31, 5.01
Coronary Heart Disease other than Angina Pectoris	1.70	0.60, 4.80	4.44	2.16, 9.10	1.53	0.72, 3.23	1.83	0.85, 3.95
Complicated Angina Pectoris	5.10	2.30, 11.31	14.00	7.46, 26.29	7.53	3.33, 17.02	4.91	2.10, 11.47
Congestive Heart Failure	9.09	3.34, 24.76	0.97	0.20, 4.58	2.07	0.97, 4.40	2.96	0.98, 8.94

TABLE 4—Association between the Disability Scale and Coronary Heart Disease Manifestations and Congestive Heart Failure by Sex and Age, from Logistic Regression Analyses (Terms in the model: the four disease categories listed below, age, and additional covariates*)

	Women		Men		Women		Men	
Disease Category	Odds Ratio	95% CI	Odds Ratio	95% CI	Odds Ratio	95% CI	Odds Ratio	95% CI
	Ages 55–69 years (N = 918)		Ages 56–69 years (N = 695)		Ages 70–88 years (N = 593)		Ages 70–86 years (N = 370)	
Uncomplicated Angina Pectoris	4.86	2.66, 8.85	5.39	2.6, 11.03	1.98	1.20, 3.27	3.28	1.62, 6.66
Coronary Heart Disease other than Angina pectoris	1.45	0.51, 4.14	3.55	1.63, 7.76	1.19	0.55, 2.58	1.76	0.78, 3.98
Complicated Angina Pectoris	5.27	2.30, 12.10	11.02	5.67, 21.42	5.92	2.48, 14.11	5.65	2.29, 13.99
Congestive Heart Failure	6.52	2.32, 18.31	0.99	0.19, 5.11	1.80	0.84, 3.86	1.79	0.56, 5.71

*Additional covariates: hypertension, cigarette smoking, body mass index, blood sugar, stroke, intermittent claudication, amputations, arthritis, back problems, gastrointestinal problems, neorological problems, and repsiratory problems.

for selection factors which could play a role in this study, factors somewhat parallel to the healthy worker effect. Participants are very possibly as a whole healthier in general than nonparticipants. As the longitudinal study progresses, widowhood, geographical relocation, and entrance into institutionalized care facilities probably account for differences in the health status of participants and nonparticipants. For example, one can surmise that the older male participants with congestive heart failure have less severe cases than the surviving cohort members who did not appear for examination 14/15; the milder nature of their disease may partially account for the lack of a relationship with disability.

The loss of severe cases through mortality is probably more relevant among the older persons. The loss of participants results in a smaller population and therefore a possible power problem in our analyses of the 70–88 year olds. It is also important to consider that as age increases, disability due to other conditions for which we have not controlled for such as depression or senility, may interfere increasingly with the assessment of CHD-related disability. Another limitation of our analyses is the absence of consideration of psychosocial variables which may influence physical functioning and consequently could confound our results.

In spite of factors which could attenuate the associations between CHD and disability, suffering from AP was a predictor of disability in each gender and age subgroup. This finding provides epidemiological support for what is commonly observed in clinical practice: once a person develops AP, it chronically affects physical function, regardless of the presence of CHD risk factors, stroke, intermittent claudication, and self-reported health conditions other than cardiovascular disease. Complicated AP was a stronger predictor of disability than uncomplicated AP. This may be explained in part by the differing severity of AP in those with and without other CHD and in part because of limited myocardial reserve imposed by myocardial infarction. That these relationships were generally stronger in men than women could reflect the fact that men suffer more severe disease than women, or that men may initially be more active, enabling their disease to have a greater potential for being more limiting.

The Framingham Study protocol dictated that once an individual was diagnosed as having AP, the individual was never re-evaluated for presence of AP. This limitation prohibited us from identifying those persons who were suffering from AP at examination 14/15; our classification of AP was essentially AP at any time from examination 1 through examination 14/15. We suspect that had we been able to use a determination of current AP sufferers, the relationship to disability would have been even stronger than our analyses indicated.

In sharp contrast to the strongly positive relationship between AP and disability, CHD other than AP was a significant predictor of disability only in the younger men. The finding is consistent with the explanations mentioned above for comparisons between older and younger participants, and between women and men. In these 40 women and 73 men diagnosed as having CHD other than AP, the mean amount of time that elapsed from the most recent event to the time of the disability measurement was greater for the older groups. This fact further helps to explain why the relation between disease and disability are weaker in the older groups; our data seem to indicate that initial disability with acute manifestations probably lessens or disappears with time. This interpretation is consistent with the results of a study of 9,385 men and women aged 18–103 years.[16] In this study a physical functioning measure, based on self-reported capacity to perform a variety of physical activities, was significantly related to a multivariate model to occurrence of myocardial infarction within the prior 12 months. Angina was also found to be significantly related to physical functioning.

The ostensibly puzzling difference between men and women in their congestive heart failure-disability relationships (significant in women but not in men) may be understood better when the relationships between CHD and congestive heart failure are examined. Among the men, 85.7 percent of those with congestive heart failure also had CHD, whereas only 47.5 percent of the women with congestive heart failure were diagnosed as also having CHD. One could hypothesize that in the 21 men with congestive heart failure, their CHD was so disabling that additional disability could not be detected. Furthermore, the very low congestive heart failure prevalence in the men probably works against the detection of a relation with disability.

While we found only one other study comparing disability in persons with and without CHD,[16] other studies have examined predictors of cardiac disability in patient-based populations, the results of which are largely complementary to ours. In a study of men ages 35–77 years with CHD, an activity score was calculated from a self-administered instrument and included such items as sports, walking, sex, and heavy housework; regression analysis on 98 of these men showed that angina was related to total activity score, but past myocardial infarction history was not so related.[6] Chirikos and Nickel studied 976 men and women (age range not identified) who were hospitalized for acute coronary attacks and found in their multivariate analyses that presence of cardiac and chronic diseases, including angina, was significantly related to restoration of functional capacity as determined by telephone interview at 6, 18 and 24 months post-hospitalization.[7] They also noted that gender was not significantly related to the restoration of functional capacity. Results from Hlatky, *et al*, are very different; in their study, a self-administered questionnaire was used to determine whether the study participants were not working because of disability.[5] Among 814 men younger than 60 years with coronary artery disease, an anginal pain scale was not related to work disability, but history of myocardial infarction was so related. This study utilized a very different type of disability measure, which might partially explain why its results differ from ours and above-mentioned studies.

In addition to the selectivity factor previously noted, another limitation of the study design warrants mention. Our study question assumes that disability developed subsequent to disease. Because assessment of physical disability status was made at a single point in time (examination 14/15), it is possible that some persons were disabled at the start of their participation in the Framingham Study or otherwise preceded the diagnosis of CHD. The items comprising the disability scale have been asked at a later Framingham examination; analyses which include these data will hopefully enable a study of change of disease status with change in disability status.

In conclusion, our analyses revealed a striking difference in the disabling effects of AP and other CHD. AP, whether complicated or not, was predictive of disability, whereas other CHD was not, except in the youngest women. This finding is in contrast to the relative roles of CHD manifestations when the outcomes of interest are further morbidity and mortality; that is, AP is less predictive of CHD morbidity and mortality than other CHD.[17] The interactions between disability and disease are complex, and it is important to keep in mind that the presence of disease may not necessarily contribute to disability, further morbidity, and mortality to a similar extent. Finally, age is another important factor to take into consideration, as our analyses indicated differences in the strength of the relationship of disability to CHD in different age groups.

These results further the cause of preventive cardiology. Our findings provide population-based evidence to back up clinical knowledge of consequences of CHD. Motivation for control of cardiovascular risk factors is not only the avoidance of well-documented morbidity and mortality from CHD other than angina pectoris; our study demonstrates clearly and provides quantification of the disabling nature of angina pectoris.

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The Effects of Psychosocial Work Organization on Patterns of Cigarette Smoking among Male Chemical Plant Employees

KATHRYN L. GREEN, SCD, AND JEFFREY V. JOHNSON, PHD

Abstract: We tested the hypothesis that job strain (the combination of high psychological job demands and low work control) is positively associated with smoking prevalence and intensity in a study group of 389 males employed in a chemical plant, using a self-administered questionnaire. In a logistic regression analysis which controlled for a number of sociodemographic factors, job strain was not found to be associated with smoking cessation. However, among smokers, those in higher-strain jobs smoked more heavily than those in lower-strain positions (OR 1.70, 95% CI = 1.10, 2.61) and were more likely to have increased the amount they smoke (OR 3.72, 95% CI = 1.92, 7.17). (*Am J Public Health* 1990; 80:1368–1371.)

Introduction

Cigarette smoking is a harmful habit, not only in terms of its health consequences but also with respect to its financial impact. Workers who smoke are said to have higher rates of absenteeism, hospitalization, illness, and job-related accidents; each employee who smokes is estimated to cost business $300 to $1000 a year.[1] Corporate concern about such facts is indicated by the current proliferation of worksite anti-smoking policies and cessation programs. However, in contrast to the extensive documentation of smoking's effects on employment, the converse of this relationship—how work may influence smoking behavior—remains largely ignored.

Overall, blue-collar workers are much more likely to smoke and less likely to quit smoking.[2] The differential is somewhat smaller for women than for men. While the proportion of persons who have made a serious attempt to quit does not differ substantially between white- and blue-collar workers (64 vs 59 percent, respectively), white-collar workers are more likely to have succeeded (48 vs 35 percent success).[3]

The relation between occupation and smoking may be mediated by the psychosocial work environment. Despite a lack of consistency in definitions and measurement, research has generally found work-related stress to inhibit cessation[4,5] and increase intensity of smoking.[4-7] Workplace social support may affect smoking behavior negatively or positively, depending on the prevalence of smoking and attitudes towards quitting among co-workers.[4,8] In addition, studies of work socialization have indicated that low job control contributes to psychological characteristics such as poor self-esteem and external locus of control,[9,10] which have been associated with a lower likelihood of quitting smoking.[11,12]

Karasek and his colleagues have defined job strain as the combination of high psychological demands with low decision-latitude or control.[13] Research in Sweden and the United States using this model has found high-strain work—where job demands are high and there is little opportunity for control—to be associated with an increased risk of cardiovascular disease (CVD), independent of smoking.[13-18] More recently, a modified version of this formulation which includes social support has identified a high CVD risk group of strained and socially isolated workers.[19,20] If job stress or strain also increases smoking, it would pose a double risk to cardiovascular health: directly, through physiological mechanisms, and indirectly, through behavioral risk factors such as smoking.

The central hypothesis of the present study was that job strain, defined as a high ratio of psychological demands to control, is negatively associated with smoking cessation and positively associated with both smoking intensity and an increase in smoking over time. A secondary objective was to explore the relationship between workplace social support and smoking behavior.

Methods

Study Group

The population surveyed consisted of the approximately 800 employees at a chemical manufacturing plant in the northeastern United States. Two years prior to the study, the plant had established an on-site "Wellness Center," consisting of exercise facilities staffed by two health educators.

Each employee was mailed a copy of the questionnaire described below; after two mailed follow-ups to non-respondents, a final response rate of 60 percent among full-time employees was obtained. From the 466 completed questionnaires, 23 females and 54 males under the age of 40 were excluded, since their work histories were substantially different from those of the majority of the workforce and their numbers too small to allow for meaningful subanalysis. Thus, all analyses were conducted on a study group of 389 male respondents ages 40 and above. The typical study group member was White (93 percent), married (85 percent), and had a high school (51 percent) or more than high school (29 percent) education. Over half of respondents reported earning between $30,000 and $40,000 annually.

Non-respondents were less likely than respondents to have registered with the exercise center (28 percent vs 39 percent), and less likely to be salaried (22 percent vs 34 percent) (as opposed to being paid hourly).

Table 1 presents the distribution of job titles in the study group, classified in seven occupational categories. Almost three-fourths were paid on an hourly basis. Half of the group, mostly hourly employees, worked shifts. The average length of time with the company was just over 25 years.

Questionnaire

The questionnaire collected information on job characteristics, health behavior, and demographics.

Address reprint requests to Kathryn L. Green, ScD, Assistant Professor, Department of Community Health & Epidemiology, University of Saskatchewan, Saskatoon, SK S7N 0W0. Dr. Johnson is Assistant Professor, Division of Behavioral Sciences and Health Education, Department of Health Policy and Management, Johns Hopkins School of Hygiene and Public Health, Baltimore. This paper, submitted to the Journal August 14, 1989, was revised and accepted for publication April 17, 1990.

TABLE 1—Distribution of Occupations in Male Study Participants

Occupational Category	N	%
Unskilled workers	26	(6.8)
Semi-skilled workers, machine operators	151	(39.2)
Skilled manual workers	87	(22.6)
Clerks, technicians	74	(19.2)
Administrative personnel, minor professionals	17	(4.4)
Business managers, lesser professionals	18	(4.7)
Higher executives, major professionals	12	(3.1)
No answer	4	
Total	389	

Job Characteristics: This section was based on Karasek's Job Content Instrument (JCI),* with minor modifications. The JCI was designed to measure the content of work tasks in a way which is applicable to all jobs in the United States. Its main focus is the psychological and social structure of work. The six job characteristic scales formed from this questionnaire assess 1) job control, 2) psychological demands, 3) physical demands, 4) job insecurity, and 5–6) social support from co-workers and from supervisors. In addition, a seventh variable measuring job strain was created by dividing the standardized score on the psychological demands scale by the standardized control score. In accord with the job strain hypothesis, the higher the demand/control ratio, the higher the level of strain. (Details of scale content and construction are in the Appendix.)

Smoking Behavior: From the questions on smoking behavior, three dichotomous smoking variables were created.

1) Smoking cessation (smokers vs ex-smokers);

2) Smoking intensity (less vs more than a pack a day);

3) Change in smoking habits since starting to work for the company (started smoking or increased amount vs no change or decreased).

Statistical Analysis

For the bivariate analyses, simple odds ratios (ORs) measured the strength of the associations, with corresponding 95 percent confidence intervals (95% CIs) calculated using Woolf's method.[21]

Multivariate analysis of the relations between job characteristics, demographic factors, and smoking behavior was conducted. In a logistic regression analysis, a stepwise backward selection procedure was used to determine which of the job characteristics and/or demographic variables were predictors of each smoking behavior variable with the inclusion level set at 0.10.

Standardized ORs (SORs) with 95 percent confidence intervals were computed from the standardized logistic coefficients, following the method proposed by Rosenmann, *et al.*[22] To indicate a negative association, a [−] sign is used, rather than the reciprocal of the SOR. The following variables were initially entered into the logistic model: job strain (demand/control ratio), job insecurity, physical demands, co-worker support, supervisor support, shiftwork (1 = work shifts; 0 = do not), marital status (1 = divorced, 0 = other), education, income, age, and race (1 = White, 0 = other).

To further examine the relation between job characteristics and behavior, multiple contingency tables were formed by stratifying the data into categories of each of the other variables retained in the logistic model, alone and in combination. From these tables, an estimate of a uniform OR over all strata was calculated using Mantel-Haenszel procedures.[23] The heterogeneity of the ORs across strata was determined by a likelihood ratio test of heterogeneity.[24] Confidence intervals around the Mantel-Haenszel point estimates were then calculated using Miettinen's test-based interval estimation procedure.[25] This analysis produced generally similar results to logistic regression, with all ORs heterogeneous across strata; thus, multiple contingency tables are discussed here only where the stratum-specific ORs are of interest.

Results

Almost half (49 percent) of the 389 study participants had quit smoking and 31 percent were current smokers. Half of the former smokers (50 percent) had quit 11 or more years earlier; just over one-third (36 percent) had quit between two and 10 years earlier and 14 percent had quit within the preceding year.

Almost half of the 117 smokers (46 percent) reported smoking more than 20 cigarettes per day; a slightly smaller proportion (41 percent) smoked 10 to 20 a day, and 13 percent smoked fewer than 10. In response to the question of whether their smoking habits had changed since starting to work for this company, 52 percent reported no change, 35 percent an increase, 10 percent a decrease, and 3 percent said they had only begun to smoke since starting to work.

Bivariate Analyses

Bivariate analyses found that, as predicted, smokers in high-strain jobs are much more likely to smoke heavily and to report an increase in smoking since starting to work for the company; those in low-strain jobs are also more likely to have quit smoking, although this association is weaker (Table 2).

There is a steady increase in the proportion of heavy smokers as job strain increases. Similarly, the proportion of smokers reporting an increase in smoking grows progressively with increasing job strain. Furthermore, the proportion of smokers who have decreased or maintained their smoking is inversely related to job strain, with a significant number reporting a decrease in smoking only at the lowest level of strain.

Multivariate Analyses

Controlling for demographic and other job characteristic variables through logistic regression reduces the association between strain and cessation (Table 3), but the relations between job strain and intensity of smoking and between job strain and increase in smoking remain strong (Tables 4 and 5). In each case, adjusting for the demographic variables remain-

TABLE 2—Bivariate Relations between Smoking Behavior Variables and Job Strain (Low vs High)

Smoking Behavior Variables	Odds Ratio (95% CI)
Cessation (former vs current smokers)	1.71 (0.92, 3.17)
Intensity (≤20 cigarettes per day vs >20)	3.86 (1.35, 10.99)
Change in smoking habits (increased/started vs no change/decreased)	7.78 (2.28, 26.50)

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*Job/Heart Project: Job Content Questionnaire and User's Guide, Revision 1.1. New York: Columbia University, March 1985. Unpublished manuscript.

TABLE 3—Smoking Cessation: Variables Retained in Logistic Model

Variables	Standardized Odds Ratio (95% CI)
Age	−1.45 (−1.14, −1.86)
Divorce	1.37 (1.07,1.75)
Co-worker Support	1.35 (1.05,1.77)
Job Strain	1.17 (0.90,1.51)

TABLE 4—Smoking Intensity: Variables Retained in Logistic Model

Variables	Standardized Odds Ratio (95% CI)
Job Strain	1.70 (1.10,2.61)
Education	−1.61 (−1.05, −2.47)
Income	1.50 (0.97,2.37)

TABLE 5—Change in Smoking Habits: Variables Retained in Logistic Model

Variables	Standardized Odds Ratio (95% CI)
Job Strain	3.72 (1.92,7.17)
Age	−2.16 (−1.32, −3.53)
Education	−2.09 (−1.22, −3.55)
Income	1.70 (1.00,2.89)

ing in the model strengthens the association between job strain and smoking behavior.

The interrelations between smoking cessation, co-worker support, and demographic variables were examined further in multiple contingency table analysis. Overall, co-worker support is negatively associated with having quit smoking. Adjusting the relation between co-worker support and quitting for both education and age reveals some interesting interactions. Co-worker support has little impact on quitting among younger men, regardless of education. Among older men with a high school education or less, co-worker support is associated with smoking (OR 3.13, 95% CI = 1.33, 6.91), while for older men with more education, it is unstably associated with quitting (OR 4.40, 95% CI = 0.82, 20.37).

Discussion

Our failure to find a relation between job strain and quitting may have been due to the cross-sectional nature of the study: half of the ex-smokers had quit more than 10 years earlier and one-fourth 20 or more years earlier, at which time their job characteristics might have been quite different. Another factor which could affect this association is awareness of occupational exposure to carcinogens. In the 1960s and '70s, a number of workers at this plant had developed lung cancer as a result of their in-plant exposure. Most of the older production workers were at the plant during that time and may have been motivated to quit smoking by the deaths of their co-workers.

As hypothesized, smokers who experience an excess of psychological demands relative to control tend to smoke more heavily and are more likely to report an increase in smoking over time than those in jobs where demands and control are better balanced. The question on which the association between job strain and increase in smoking is based relied on self-reported assessment of change over a long period of time. Smokers had worked for the company from nine to 42 years, half between 21 and 29 years. However, the accuracy of the report is suggested by the fact that those who reported an increase in smoking were much more likely to smoke over a pack a day (OR 2.81, 95% CI = 1.24, 5.97).

The most important occupational predictor of smoking cessation in this group was co-worker support. While the overall relation is not strong, the stratified analysis by age and education indicates statistical interactions. For older smokers with a high school education or less, strong co-worker support was associated with a lower likelihood of having quit smoking. For those with more education, the opposite is true: those with strong social support were *more* likely to have quit.

The finding that the effect of social support depends on educational level is consistent with the fact that in this study group, smoking was more prevalent among less-educated older men than among those with more education (31 vs 9 percent current smokers, respectively) and that the former were more likely to report that most of the people with whom they spend time do not care whether they smoke, while the latter tended to believe that most people would like them to quit. Other studies also suggest that higher education is associated with more negative attitudes toward smoking.[26] However, a plausible alternative explanation for this finding is that smoking status affects co-worker support: those who work with a group of smokers may fit in better if they also smoke, while being one of a minority of smokers could impair relations with non-smoking co-workers.

If job strain does increase the intensity of smoking, attempts to encourage smokers in high-strain jobs to quit smoking may be hampered by the nature of their work. Thus, modifying job structure so as to increase control and thereby decrease strain could enhance the success of cessation efforts in this population. This recommendation may seem unnecessary and inappropriate if smoking is simply considered a personal behavior controlled by the individual. However, as Syme has pointed out, the systematic patterning of smoking rates in society strongly suggests that social and cultural forces play an important role in this behavior.[26]

APPENDIX

Construction of Job Characteristic Scales

Control

This scale sums responses to 15 items: six items measure skill discretion (degree to which the job involves learning new things, repetitiveness, creativity, varied tasks, and development of the individual's special abilities), seven items measure decision authority (individual's ability to make decisions about his/her own job, to influence the work group, and to influence company policy), and two items measure closeness of supervision (alpha = .85).

Psychological Demands

This scale sums responses to 10 items assessing how fast and how hard the respondent has to work, hectic nature of the job, whether there is enough time to get the job done, frequency of work under time pressure, amount of work and level of concentration required, presence of conflicting demands, and how often tasks are interrupted or work is slowed down by waiting for other people or departments (alpha = .78).

Job Insecurity

This scale sums responses to six items: level of job security, prospects for career development and promotion, future value of skills, possibility of lay-off in recent past and near future, and ability to keep up with work (alpha = .69).

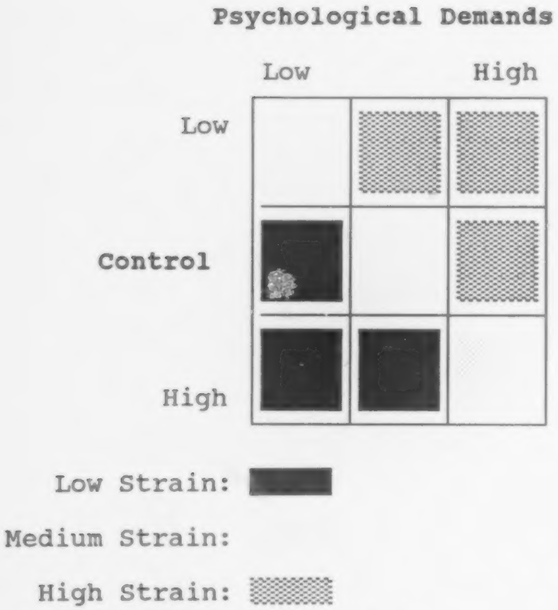

FIGURE 1A—Construction of Three-Level Strain Variable

Physical Demands

This scale sums responses to five items measuring physical exertion (degree to which job involves fast and continuous physical activity, intensive physical effort, and lifting) and ergonomic strain (how often body or head and arms must be held in awkward position) (alpha = .91).

Co-worker Social Support

This scale sums responses to four items measuring co-workers' competence, helpfulness in getting the job done, friendliness, and degree to which they take a personal interest in the respondent (alpha = .71).

Supervisor Social Support

This scale sums five items measuring the supervisor's concern for subordinates' welfare, competence, helpfulness in getting the job done, hostility, and degree to which he/she pays attention to what respondent says (alpha = .84).

Each of these scales was standardized by subtracting the mean from the raw score and dividing the result by the standard deviation; a constant was then added to avoid negative scores. To form contingency tables for bivariate and multivariate analyses, these standardized scores were divided into approximate tertiles (except for the two social support variables which could only be divided roughly at the median, because of their distribution).

Job Strain

In addition to the continuous job strain measure (demand/control ratio), a three-level ordinal form was created from a nine-cell matrix constructed from the three levels of control and psychological demands. Each of these nine cells was then classified into one of three levels of strain, based on the Demand-Control Model, which hypothesizes a dimension of strain running diagonally from the lowest level of demands and highest level of control to the opposite extreme (Figure 1A).

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Use of Fat-Modified Food Products to Change Dietary Fat Intake of Young People

R. Curtis Ellison, MD, Robert J. Goldberg, PhD, Jelia C. Witschi, MS, RD, Ann L. Capper, RD, Elaine M. Puleo, PhD, and Fredrick J. Stare, MD, PhD

Abstract: Food purchasing and preparation practices were modified in two boarding high schools to increase the polyunsaturated-to-saturated fat ratio (P/S) of the diet of students by changing food products rather than attempting to change eating behaviors. During years when fat-modified products were served, the P/S of males increased by 75 percent, versus a decrease of 6 percent during control years. For females, P/S increased by 53 percent during intervention years, versus an increase of 6 percent during control years. (*Am J Public Health* 1990; 80:1374–1376.)

Introduction

In an effort primarily to decrease the risk of premature cardiovascular disease, many health promotion organizations currently advise that all Americans, including young people, modify the amount and type of fat in their diets.[1-3] However, the traditional, one-on-one counseling approach to changing eating behaviors is difficult and time-consuming and is an inefficient method for changing the dietary practices of an entire population. An alternate approach is to change the fat content of food products. This would appear to be particularly relevant today, since such a large proportion of foods consumed in the United States are manufactured, processed, or prepared by the industry, rather than prepared from scratch in the home. This limits the control that individuals have over the amount and type of fat that is contained in the foods they consume.

Studies were carried out at two boarding high schools in New England to determine the impact of modifying fats in food products served in the schools' dining halls on the polyunsaturated-to-saturated fat ratio (P/S) of the diets of students. The goal was to determine the effects on the total intake of nutrients of a "passive" intervention, i.e. an intervention at the level of food production and preparation only. Foods with a higher P/S ratio were purchased and changes were made in food preparation practices in the schools' kitchens, but students were not advised to modify their eating behaviors. This paper reports the effects on students' overall dietary fat intake when the fat-modified products were introduced into the dining halls of each of the two schools over a school year.

Methods

The project was carried out at Phillips Exeter Academy in Exeter, New Hampshire and Phillips Academy in Andover, Massachusetts during 1986–88. School food service workers at each institution were trained to procure and prepare food products in which some of the saturated fat content was replaced by polyunsaturated fat.[4,5] It was sought to avoid alteration of the total fat content of food products. The general intent was to make only those modifications that were feasible in terms of product availability, acceptability, and cost, at the time of the study, for an institutional food service operation.

The study design was that of a longitudinal investigation with food product modification being carried out in each of two schools in alternate school years. After the first month of each school year, modifications of food purchasing and preparation practices were carried out in the school which was serving as the intervention school for the year, while the other school continued to serve usual food products. Some modification was carried out for almost two-thirds of fat-containing foods served. While students were aware that some foods were being prepared using different types of fat, they were not told and, in general, could not tell which items had been modified.

For each of the two years, all students at both schools who were taking selected courses in biology or chemistry (approximately 200 at each school each year) were trained and asked to furnish food diaries at specified times during the year to monitor the effects of the food service changes on their intakes of calories and specific nutrients. All foods ingested were recorded, including those foods that were obtained outside of the dining halls (which furnished 28 percent of calories).[6] Comparisons were made between trends in fat intake in the two schools between a baseline period in the fall (when both schools were serving usual products) and a follow-up period in the spring, when the intervention school students were being served fat-modified products. Analysis of food diaries was performed by trained registered dietitians using the "Food Finder" program at the Center for Clinical Computing, Beth Israel Hospital, Boston, MA.[7]

A total of 6,436 food diaries were submitted during the control and intervention years at the two academies. Data provided by all students completing food diaries during a period were used to calculate mean values for that period.

Results

Baseline and follow-up dietary data for males and females are given in Tables 1 and 2, respectively. At baseline, all groups were consuming approximately 35 percent of their calories from fat. During intervention years, the P/S ratio of the students' diet increased by 75 percent (from 0.53 to 0.93) for males and by 53 percent (from 0.64 to 0.98) for females. During control years, the P/S ratio decreased by 6 percent for males and increased by 6 percent for females.

Discussion

The present study was designed to test the feasibility of changing the P/S ratio of the diet of boarding school students

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From the Section of Preventive Medicine and Epidemiology, Evans Department of Medicine, Boston University School of Medicine; the Department of Medicine, University of Massachusetts Medical School; the Department of Nutrition, Harvard School of Public Health; and the Division of Public Health, University of Massachusetts, Amherst. Address reprint requests to R. Curtis Ellison, MD, University Hospital, Room B-612, 88 E. Newton Street, Boston, MA 02118. This paper, submitted to the Journal September 29, 1989, was revised and accepted for publication February 28, 1990.

TABLE 1—Baseline and Follow-up Dietary Data among Males, by Intervention Status (Mean ± SD)

	Control			Intervention			Intervention Effect (Intervention-Control)	
	Baseline (n = 197)	Follow-up (n = 109)	Diff. (%)	Baseline (n = 198)	Follow-up (n = 98)	Diff. (%)	Effect (%)	(95% CI)
Calories	2831 ± 801	2603 ± 786	−8	2835 ± 802	2420 ± 688	−15	−7	(−14, +1)
Total Fat (%kcal)	35.5 ± 5.0	33.6 ± 5.6	−5	34.5 ± 5.4	31.9 ± 6.5	−8	−2	(−7, +3)
Total Fat (gm)	112.1 ± 37.7	98.7 ± 34.5	−12	110.9 ± 41.3	86.6 ± 30.3	−22	−10	(−19, −1)
Polyunsaturated fat (%kcal)	6.1 ± 1.8	5.5 ± 1.7	−10	5.7 ± 1.8	7.9 ± 3.4	+39	+48	(+34, +63)
Polyunsaturated fat (gm)	19.3 ± 7.6	16.2 ± 6.9	−16	17.8 ± 7.4	21.1 ± 9.2	+19	+35	(+21, +48)
Monounsaturated fat (%kcal)	12.5 ± 2.4	12.1 ± 2.4	−3	12.5 ± 2.4	10.9 ± 2.6	−13	−10	(−15, −4)
Monounsaturated fat (gm)	39.8 ± 14.4	35.6 ± 13.1	−11	40.5 ± 16.9	29.7 ± 11.5	−27	−16	(−25, −8)
Saturated fat (%kcal)	12.3 ± 2.6	12.0 ± 2.6	−2	12.2 ± 2.6	9.6 ± 2.8	−21	−19	(−25, −13)
Saturated fat (gm)	39.5 ± 14.9	35.1 ± 13.1	−11	37.8 ± 15.6	26.0 ± 11.5	−31	−20	(−29, −11)
P/S Ratio	0.54 ± 0.20	0.51 ± 0.22	−6	0.53 ± 0.21	0.93 ± 0.45	+75	+81	(+60, +102)

TABLE 2—Baseline and Follow-up Dietary Data among Females, by Intervention Status (Mean ± SD)

	Control			Intervention			Intervention Effect (Intervention-Control)	
	Baseline (n = 188)	Follow-up (n = 130)	Diff. (%)	Baseline (n = 191)	Follow-up (n = 130)	Diff. (%)	Effect (%)	(95% CI)
Calories	2040 ± 552	1898 ± 649	−7	2063 ± 519	1773 ± 581	−14	−7	(−14, −0)
Total Fat (% kcal)	34.9 ± 5.5	34.1 ± 6.0	−2	34.2 ± 5.9	30.7 ± 6.3	−10	−8	(−12, −3)
Total Fat (gm)	80.9 ± 27.2	73.7 ± 31.6	−9	81.1 ± 27.2	62.1 ± 26.9	−23	−15	(−23, −6)
Polyunsaturated fat (%kcal)	6.5 ± 2.3	6.5 ± 2.6	0	6.4 ± 2.0	7.8 ± 2.9	+22	+22	(+11, +33)
Polyunsaturated fat (gm)	14.5 ± 5.8	13.2 ± 5.1	−9	14.5 ± 5.9	15.7 ± 7.6	+8	+17	(+6, +29)
Monounsaturated fat (%kcal)	11.7 ± 2.5	11.3 ± 2.6	−3	11.6 ± 2.6	9.8 ± 2.5	−16	−12	(−18, −6)
Monounsaturated fat (gm)	27.2 ± 10.3	24.6 ± 11.8	−10	27.5 ± 10.3	20.1 ± 10.3	−27	−17	(−27, −8)
Saturated fat (% kcal)	12.1 ± 2.9	11.7 ± 3.3	−3	11.8 ± 2.8	9.2 ± 2.7	−22	−19	(−25, −12)
Saturated fat (gm)	28.0 ± 10.7	25.9 ± 13.4	−8	26.6 ± 10.2	18.6 ± 9.7	−30	−23	(−33, −13)
P/S Ratio	0.62 ± 0.28	0.66 ± 0.40	+6	0.64 ± 0.25	0.98 ± 0.42	+53	+47	(+31, +62)

by modifying the foods served rather than by changing student eating behaviors. The modified products were very well accepted by students. While the estimated effect of the intervention on the diets of students is based on their self-reports of intake, the students had not been told to change their dietary habits or fat intake and thus were unlikely to be trying to please the investigators by reporting what was believed to be desired. Also, in general, they did not know which products had been modified. The study demonstrated that the modification of approximately two-thirds of the fat-containing food products served in the schools' dining halls reduced total saturated fat intake by over 20 percent and resulted in a P/S ratio approaching 1.0, the level currently recommended by most health agencies.[1-3]

The approaches tested in this study are commercially feasible and could be implemented in schools or other institutional food service operations elsewhere. Furthermore, the approaches are applicable to food manufacturers and processors in general. The development of margarines with higher P/S ratios and the replacement of highly saturated coconut, palm, or palm kernel oils with less saturated oils in many food products are examples of such applications. The widespread adoption of such approaches by the food industry could result in healthier foods for Americans.

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Weekly Food Servings and Participation in Social Programs among Low Income Families

Douglas L. Taren, PhD, Warren Clark, DMin, MDiv, Mary Chernesky, MS, CHE, and Elisabeth Quirk, MEd

Abstract: Low income families were interviewed to determine factors related to the number of family food servings per week. A multiple regression model indicated that participation in WIC (supplemental food program for women, infants and children), household size, and number of different income sources were associated with more family food servings per week. Number of food servings per week decreased the last week of the month most often in families with younger members. (*Am J Public Health* 1990; 80:1376–1378.)

Introduction

Several studies suggest that family food shortages have increased during the past decade. Populations in the United States that are at risk for these shortages include the elderly, single parent families, and children.[1–10] Other studies have shown that participants of government assistance programs also endure food shortages.[11–16] These food shortages follow monthly cycles with the greatest prevalence occurring the last week of the month.[12]

The purpose of this study was to determine if the number of food servings per week could be used to assess food shortages that occur at the end of the month, the number and types of food and income assistance programs used by low income families, and factors associated with the number of family food servings.

Methods

A random sample of 25 percent of family members from six food cooperatives for low income families was selected, with 109 of 114 responding. Fifty families from the Hillsborough County Florida Expanded Food and Nutrition Education Program (EFNEP) who were not participating in the food cooperative were randomly selected to participate in the study; 12 did not participate and 10 more were randomly selected as substitutes. One home interview was conducted for each family.

The focus of the study was the household. Demographic data were collected with information on participation in 23 food and income assistance programs. The number of family food servings during the past week from a list of 27 food groups was collected. A serving was defined as the preparation and offering of a food group to family members; second and third offerings of the food during the same meal were not counted. A normally distributed additive score for 20 food groups was calculated (mean ± 1 SD: 24.7 ± 5.0; Cronbach's alpha of 0.76.)[17] The excluded food groups were oils, salty snacks, candy, sweet baked goods, soda, soup, and coffee.

Frequency analysis using Chi-square and T-tests for means were used to determine the characteristics of the participants of the two programs. Multiple regression was used to determine factors related to the number of food servings. Potential independent variables for the final model were: participation in individual food and income assistance programs; the number of income or food assistance programs; the reference week of the month for the servings; the household size; sex of respondent; the age of the household respondent; participation in the food cooperatives or EFNEP; and an interaction term between program participation and reference week. However, when the final regression model was constructed several variables were not significant and dropped.

Results

The EFNEP households compared with the food cooperative families were younger: (31 vs 51 years); had more single parents (60 percent vs 40 percent); had more children less than 5 years of age (60 percent vs 44 percent); and had lived less time in their present location (42 vs 128 months).

The food and income sources used by these families are presented in Table 1. There was a wide variation in the combination of sources used among all families with no two families using the same combination; 39 percent had four or more sources and 13 percent had six or seven sources of income.

The last week of the month was associated with a decrease in the number of food servings per week (mean ± 1 SEM 22.9 ± 0.8) compared with the second (26.4 ± 1.0) and third (26.1 ± 0.8) week of the month (Figure 1). The variables in the best model (Table 2) to predict the number of food

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Address reprint requests to Douglas L. Taren, PhD, Assistant Professor, Department of Community and Family Health, College of Public Health, University of South Florida, 3500 E. Fletcher Avenue, Suite 106, Tampa, FL 33618. Dr. Clark and Ms. Quirk are with Hunger Action, Tampa United Methodist Centers; Ms. Chernesky is with the Hillsborough County Expanded Food and Nutrition Education Program. This paper, submitted to the Journal November 13, 1989, was revised and accepted for publication May 31, 1990.

TABLE 1—Food and Income Sources That at Least 10 Percent of the Sample Used[a]

	Food Coop N = 109 %	EFNEP N = 48 %
Food Sources		
Food Pantry	13	10
Commodity	74	33
School Breakfast	32	27
School Lunch	32	33
Food Stamps	48	54
WIC	6	21
Income Sources		
AFDC	24	44
SSI	32	10
Social Security	46	10
Wages	15	27

a) Programs that had less than 10% of participating families included soup kitchens, shelters, senior centers, meals on wheels, credit from stores, retirement funds, veterans, and loans.

TABLE 2—Regression Model to Predict Number of Meals Served in Household during Past Week

	Regression Coefficient	Standard Error
Intercept = 21.7		
WIC[a]	2.3	1.2
Household Size[b]	0.7	0.2
No. Income Sources[c]	1.2	0.4
Week of Month[d]	−0.7	0.4
Week Program[e] Interaction	−0.6	0.1

$F = 14.4$, $r^2 = 0.33$

a) A value of 1 was given to families who participated and a score of 0 for families who did not participate.

b) Household size was a continuous variable.

c) Number of income sources was a continuous variable.

d) Week of the month was a categorical variable that represented the reference week of the month (first to fourth) for the food servings.

e) The interaction term between families from two programs (food cooperative and EFNEP) and the reference week of the month for food servings.

servings per week were: participation in WIC, number of income assistance programs, number of people in the household, week of the month, and an interaction between the week of the month and participation in either the food cooperatives or EFNEP. This interaction represents a greater decrease in the number of food servings the last week of the month in the EFNEP families compared with the food cooperative families.

Discussion

The results from this study support previous reports that end of the month food shortages occur in low income families.[12,18-20] The significant interaction term for the number of food servings identified the EFNEP families who were younger, with children, and less established at their current location as one group at particular risk.

Our study also provides evidence that participation in a greater number of income assistance programs was associated with food servings. However, even after controlling for these factors there was a significant decrease in the number of food servings at the end of the month. This implies that current government support may not be sufficient to last an entire month, especially when the assistance is given one time at the beginning of each month. Several studies of food stamps recipients reported that families use their monthly allotment before the end of the month.* Even families who participate in five different programs have been reported not able to meet the thrifty food plan.[20] Nevertheless, even though there is no guarantee that the diets of families participating in federal programs are of high nutritional quality,[21] assistance programs have been shown to improve the diets of low income families and improve food security.[22]

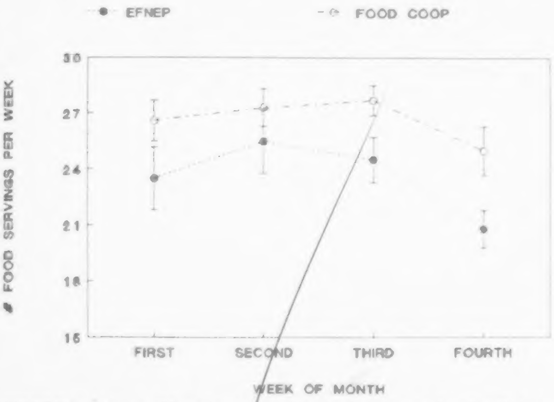

FIGURE 1—Number of Food Servings per Week in Low Income Families (Values represent the adjusted means for the number of servings per week; the vertical bars represent ± 1 standard error.)

In conclusion, weekly food servings decreased the last week of the month for these families and could be measured using a food frequency questionnaire. This pattern may be related to insufficient funds that occur the last week of the month when financial assistance from government sources are depleted.

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The Buddy Volunteer Commitment in AIDS Care

PRISCILLA VELENTGAS, CHRISTIAN BYNUM, AND SALLY ZIERLER, DRPH

Abstract: Buddy volunteers provide crucial assistance to people with HIV-related illnesses. Based on volunteers' self-administered questionnaires, our study describes the nature of buddy work. Volunteers indicated their satisfaction with both personal performance and buddy program administration. Several factors were associated with volunteer satisfaction. This report is a first attempt to describe this special relationship created in response to the human immunodeficiency virus (HIV) epidemic. (*Am J Public Health* 1990; 80:1378–1380.)

Introduction

Since the onset of the AIDS (acquired immunodeficiency syndrome) epidemic a decade ago, community-based support for people living with AIDS has increased. Formation of community-based organizations to help people with AIDS (PWAs) has taken place across the United States. Services provided range from food distribution to information hotlines to legal and political advocacy. The size of these organizations ranges from a few dozen active volunteers to several hundred.

This report focuses on programs that establish a special relationship between volunteer and client. "Buddy programs" pair a volunteer from the community with a PWA. The buddy volunteer gives his/her client companionship and basic assistance in living with AIDS, and performs tasks ranging from housekeeping to providing emotional support.[1,2] Thousands of volunteers across America have entered into this profound commitment which often lasts through the death of the person with AIDS.[3]

We attempted to investigate the work done by volunteers at one intermediate-sized AIDS project to determine the nature of the contact existing between the volunteers and the project and between the volunteers and their PWA clients.

Methods

Study Populations

To be eligible for this study, women and men must have completed a buddy training program offered by Rhode Island Project/AIDS prior to February 1989. This process included completion of an application and participation in a multiphase training program. One hundred forty-three past and present buddy volunteers were eligible for the study. To be in the study, active and former volunteers returned a self-administered questionnaire. This report is based on information from the 67 volunteers who completed the questionnaire.

Data Collection

An anonymous, self-administered questionnaire was distributed to all active and inactive buddy volunteers. Initially, questionnaires were coded to be matched with the volunteers' initial applications to the program so comparisons could be made between questionnaire and application data. Buddy support group leaders distributed the questionnaires. Inactive volunteers were mailed a questionnaire, along with a cover letter and a self-addressed stamped envelope, to their last known address.

To improve response, a second batch of questionnaires was distributed to group leaders with instructions to mail them back directly to the investigators. Inactive volunteers

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From the Department of Community Health, Division of Biology and Medicine, Brown University. Address reprint requests to Sally Zierler, DrPH, Assistant Professor of Medical Science, Box G-A405, Department of Community Health, Brown University, Providence, RI 02912. This paper, submitted to the Journal September 15, 1989, was revised and accepted for publication March 7, 1990.

who had not previously returned a questionnaire were sent another one with a second cover letter.

Data Analysis

Active and inactive populations were examined together. Satisfaction with various aspects of program administration and with personal performance as a buddy volunteer was the conceptual outcome of interest. A satisfaction score was devised to classify respondents as satisfied, dissatisfied, or indeterminate. To quantify an individual's satisfaction, we created a five-point scale relating to a volunteer's levels of satisfaction with program administration and personal performance as a buddy; among these items were queries into the nature and perception of typical requests clients made to buddy volunteers, such as help with cleaning, transportation, or meal preparation. Other questionnaire items were examined in relation to this scaled outcome.

Most differences between the two groups were observed through the comparison of proportions. Satisfaction prevalence ratios were examined for certain determinants, namely lag time between becoming a buddy and being assigned a client, number of the client's total requests for specific buddy services per week, and volunteer's assessment of client's supportive relationships. Confidence intervals for these measures were calculated using Miettinen's test-based procedure.[4]

To evaluate whether the segment of the buddy volunteer population responding to the questionnaire was representative of the buddy volunteer population as a whole, data from the applications of buddies known to have responded to the questionnaire (in the first mailing) were compared to data collected from applications of the nonrespondents.

Results

Volunteer and Client Profiles

Fifty-three percent of buddies responding to the questionnaire were male, and most volunteers (73 percent) were 26 to 45 years old. Of 61 buddies identifying with a specified race category, only five (8 percent) reported themselves as other than White. When compared with AIDS cases in Rhode Island or with the client population, minorities were underrepresented among the volunteers. According to Rhode Island Project/AIDS, their buddy program serves 63 percent of the state's IV (intravenous)-drug-related AIDS cases.

The Buddy Experience

Tables 1 and 2 show distributions of experiences as buddy volunteers: Table 2 focuses on the weekly requests that are made to volunteers from their clients; Table 1 shows most buddies responding had been assigned at least one client; over 20 percent had been assigned two or more. Over one-third of buddies had experienced the death of a client. Few buddies had requested reassignment while their clients were alive. Over half of buddies reported that being a buddy had increased the stress in their lives.

The most significant problems faced in working with PWAs were the time commitment and emotional investment (Table 1). Dissatisfied buddies reported more requests for certain kinds of tasks than satisfied volunteers, such as cleaning and preparing meals (Table 2). Almost all the buddy volunteers visited their client at least once weekly. Over half provided transportation at least once per week. Shopping, cleaning, and preparing meals were done less often. Table 3 shows a distribution of the hours spent per week by buddy volunteers according to satisfaction status. Over 60 percent of volunteers spent six or more hours weekly providing buddy services.

TABLE 1—Distributions of Factors in Buddy Experience According to Satisfaction Status [number reporting factor/number reporting satisfaction status (percent)*]

	Satisfied N (%)	Dissatisfied N (%)	Overall*** N (%)
Total Clients			
none	1/20 (5)	6/36 (17)	12/65 (18)
one	12/20 (60)	23/36 (64)	39/65 (60)
two	6/20 (30)	6/36 (17)	12/65 (18)
three	1/20 (5)	1/36 (3)	2/65 (3)
Total Deceased Clients			
none	12/20 (60)	23/36 (64)	41/64 (64)
one	6/20 (30)	10/36 (28)	18/64 (28)
two	2/20 (5)	3/36 (8)	5/64 (8)
Requests for Reassignment	1/20 (5)	3/36 (8)	6/66 (9)
Major Problem Cited			
time commitment	**	12/36 (33)	22/55 (40)
emotional investment	**	8/36 (22)	15/55 (27)
other	**	12/36 (33)	18/55 (33)
Stress Change Since Becoming a Buddy			
increased	12/20 (60)	18/33 (55)	36/66 (55)
decreased	3/20 (15)	4/33 (12)	17/66 (28)
same	5/20 (25)	11/33 (33)	13/66 (20)

*All percents are rounded to the nearest whole number; totals may not equal 100%.

**None of the "satisfied" respondents reported a major problem.

***"Overall" includes all respondents to the question: "satisfied," "dissatisfied," and intermediate.

TABLE 2—Frequency of Weekly Requests According to Satisfaction Status [number reporting request/number of respondents in satisfaction category (percent)*]

Number of Requests/Week	Satisfied N (%)	Dissatisfied N (%)	Overall** N (%)
Personal Visits			
0	1/19 (5)	1/28 (3)	2/52 (4)
1	7/19 (37)	8/29 (28)	18/52 (35)
2 or more	11/19 (58)	20/29 (69)	32/52 (62)
Shopping			
0	11/19 (58)	15/29 (52)	44/49 (90)
1	6/19 (32)	10/29 (35)	17/51 (33)
2 or more	2/19 (11)	4/29 (14)	6/51 (12)
Cleaning			
0	17/18 (95)	24/28 (86)	44/49 (90)
1	1/18 (6)	3/38 (11)	5/49 (10)
2 or more	0/18 (0)	1/28 (4)	1/49 (2)
Preparing Meals			
0	16/18 (89)	22/28 (79)	41/49 (90)
1	2/18 (11)	3/28 (11)	5/49 (10)
2 or more	0/18 (0)	3/28 (11)	3/49 (6)
Transportation			
0	7/19 (37)	13/29 (45)	22/52 (42)
1	2/28 (11)	7/29 (24)	16/52 (31)
2 or more	4/19 (21)	9/29 (31)	14/52 (27)
Telephone Calls			
0	2/18 (11)	3/29 (10)	6/51 (12)
1	5/18 (28)	7/29 (24)	13/51 (26)
2 or more	11/18 (61)	19/29 (66)	32/51 (63)

*All percents are rounded to the nearest whole number; totals may not equal 100%.

**"Overall" includes all respondents to the question: "satisfied," "dissatisfied," and intermediate.

Buddy volunteers were twice as likely to report feeling satisfied if assigned to their first PWA one to five months after acceptance to the Buddy Program at Rhode Island Project/AIDS (prevalence ratio = 2.0; 95% confidence interval = 0.8,

TABLE 3—Distribution of Total Hours Devoted to Buddy Work by Satisfaction Status [number in category/number of respondents (percent)*]

# Hours per Week Devoted	Satisfied N (%)	Dissatisfied N (%)	Total N (%)
≤5 hours/week	7/19 (36)	12/31 (39)	19/50 (38)
6–10 hours/week	6/19 (32)	6/31 (19)	12/50 (24)
≥11 hours/week	6/19 (32)	13/31 (42)	19/50 (38)

*All percents are rounded to the nearest whole number; totals may not equal 100%.

5.1). Volunteers with 10 or less requests per week from their clients also reported greater satisfaction (prevalence ratio = 3.2; 95% ci = 1.5, 6.0). Buddies serving PWAs with more supportive relationships (other than buddy-client) in their lives were twice as likely to be satisfied in their work (prevalence ratio = 2.0; 95% ci = 0.8, 4.7).

Discussion

This study may be the first attempt to describe the work of AIDS volunteer programs such as the buddy program of Rhode Island Project/AIDS. Comparison of buddy/client and client/state AIDS case demographics suggest that minorities are underrepresented in these volunteer populations. While none of the volunteers responding to our questionnaire reported themselves as representing the Latino population, Rhode Island Project/AIDS' proportion of Black and Latino clients (25 percent)[5] is comparable with the state's proportion of Black and Latino cases (29 percent).[6]

This study describes the variety of services performed by buddy volunteers, and their interaction with the agency. Although there was considerable comparability between respondents and nonrespondents with respect to motivation for volunteering and providing support as buddies, we are cautious about generalizing the survey findings because the response rate was low.

The Buddy Program is a valuable, volunteer-based health care service that is critically needed by people with HIV (human immunodeficiency virus)-related illnesses. This description may stimulate increased appreciation of this service as well as promote an understanding of how to organize volunteers' efforts and support their needs.

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Epidemiology of Reported Cases of AIDS in Lesbians, United States 1980–89

Susan Y. Chu, PhD, James W. Buehler, MD, Patricia L. Fleming, PhD, and Ruth L. Berkelman, MD

Abstract: National surveillance data for reported cases of acquired immunodeficiency syndrome (AIDS) were used to assess demographic characteristics and behavioral risk factors in lesbians. From June 1, 1980 through September 30, 1989, 79 women with AIDS reported sex relations only with a female partner; most of these women (95 percent) were intravenous drug users. Prevention of human immunodeficiency virus (HIV) infection in the lesbian community will require efforts to prevent and reduce intravenous drug use. (*Am J Public Health* 1990; 80:1380–1381.)

Introduction

Very little has been written on the epidemiology of acquired immunodeficiency syndrome (AIDS) and human immunodeficiency virus (HIV) infection in the lesbian population. Two instances of female-to-female sexual transmission of HIV have been reported,[1,2] and there is some concern that the risk of infection in female homosexuals may be underestimated because of the lack of information about this population group.[3] This report describes various demographic characteristics and behavioral risk factors of lesbians with AIDS in the United States.

Methods

We used national surveillance data for 9,717 cases of AIDS in adult women reported to the Centers for Disease Control (CDC) between June 1, 1980 and September 30, 1989. Only those that met the CDC surveillance case definition for AIDS were included in our analysis.[4] AIDS patients were grouped according to their reported sexual behavior, using responses to the following two questions:

- "After 1977 and preceding the diagnosis of AIDS, did this patient have sexual relations with a male partner?"
- "After 1977 and preceding the diagnosis of AIDS, did this patient have sexual relations with a female partner?"

From the Division of HIV/AIDS, Center for Infectious Diseases, Centers for Disease Control, Atlanta. Address reprint requests to Susan Y. Chu, PhD, HIV/AIDS, Center for Infectious Diseases, GZ9, CDC, Atlanta, GA 30333. This paper, submitted to the Journal December 27, 1989, was revised and accepted for publication March 26, 1990.

Women who had sexual relations only with female partners were classified as lesbian, women who had both male and female sexual partners were classified as bisexual, and women who had only male sexual partners were classified as heterosexual. For women who were classified as lesbian, recorded data (e.g., sex, sex partners, mode of transmission) were confirmed directly with the reporting health departments to minimize misclassification and data entry errors. We excluded 1,242 women who could not be classified as lesbian, bisexual, or heterosexual because of incomplete responses to the above two questions. However, 90 percent of these women had a known risk for HIV exposure (intravenous drug use, heterosexual contact with a person at increased risk for, or known to be, infected with HIV, or blood transfusion).

Persons with AIDS were classified in a hierarchy of exposure categories according to their presumed means of acquiring HIV infection. Those with more than one possible means of acquisition were classified only in the exposure listed first in the hierarchy, by using the following sequence: 1) intravenous drug use, 2) receipt of blood products, 3) heterosexual contact with an HIV-infected partner (or having a specified risk for HIV), and 4) "no identified risk." The last category represents persons who had none of the other risk exposures above, and therefore could include women who acquired HIV infection through female-to-female sexual contact.

Results

As of September 30, 1989, there were 79 reported cases of AIDS in lesbians, representing 0.8 percent of all reported adult women with AIDS in the United States.

As shown in Table 1, most of the lesbians with AIDS (95 percent) were intravenous drug users (IVDUs). The remaining 5 percent acquired HIV infection through the receipt of blood or blood products. As with IVDU-associated AIDS in all women, most lesbian women with AIDS were younger (85 percent were less than 40 years old), Black or Hispanic (80 percent), and resided in the northeast (77 percent). There were 103 reported cases of AIDS in bisexual women. Most bisexual women with AIDS also were IVDUs (79 percent), 16 percent had sex partners at increased risk or known to be infected with HIV, and 4 percent had histories of blood transfusion.

We also compared the occurrence of reported opportunistic diseases among homosexual, bisexual, and heterosexual women with AIDS. The only notable difference was in the occurrence of Kaposi's sarcoma (KS). KS (including presumptive and confirmed diagnoses) was not reported in any of the homosexual females with AIDS but was reported in 9 percent (9/103) of bisexual females and 2 percent (167/8,293) of heterosexual females with AIDS.

TABLE 1—AIDS cases in Lesbians (June 1980-September 1989) reported in the United States

AIDS Cases	No. (%)
Exposure category	
IVDU	75 (95)
Transfusion	4 (5)
Age (years)	
13–19	0 (0)
20–29	19 (24)
30–39	48 (61)
≥40	12 (15)
Race/ethnicity	
White	15 (19)
Black	30 (38)
Hispanic	33 (42)
Other	1 (1)
Region	
Northeast	61 (77)
Midwest	0 (0)
South	7 (9)
West	3 (4)
Territories	8 (10)

Discussion

Lesbians represent a very small proportion of women with AIDS in the United States, and most cases are related to intravenous drug use. Prevention of HIV infection in the lesbian community will require efforts to prevent and reduce IV-drug use—the major and most direct means of transmission in this population.

Although there is no specific exposure category for female-to-female transmission, lesbians who acquired HIV infection from female-to-female contact would have been classified in the "no identified risk" (NIR) exposure category, of which there were none. Female-to-female transmission of HIV also could occur among bisexual females; however, only one bisexual female had no identifiable risk and was classified in the NIR category.

Although it is not possible to determine whether differences in the occurrence of Kaposi's sarcoma (KS) reflect differences in exposures, we also observed that among women with AIDS who reported sex with a *bisexual* man, 4.1 percent (20/493) had KS reported, whereas among women who did not report sex with a *bisexual* man, 1.6 percent (52/3162) reported KS. Moreover, the frequencies of reported KS were higher among homosexual (22 percent) or bisexual men (14 percent) than among heterosexual men (3 percent) (whose predominant means of HIV exposure was intravenous drug use). These patterns are compatible with the hypothesis that KS is a sexually transmitted infection that occurs in association with HIV infection and disproportionately affects homosexual and bisexual men.[5]

It is important to note that we used categories based on reported sexual behaviors, not on self-labeled sexual orientation. The labels "heterosexual" and "homosexual" do not always predict sexual behavior.[6] Our "lesbian" category does not necessarily represent women who identify themselves as lesbians, but more accurately represents women with AIDS who reported having sexual relations only with a female partner since 1977.

Although female-to-female transmission of HIV appears to be an extremely rare event, the occurrence of AIDS among lesbian and bisexual women indicates that women who engage in sex with other women can be exposed to HIV. Case reports of female-to-female sexual transmission of HIV[1,2] and the well-documented risk of female-to-male HIV transmission indicate that vaginal secretions and menstrual blood are potentially infectious and that mucus membrane (e.g., oral, vaginal) exposure to these secretions can potentially lead to HIV transmission.

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Prior Cesarean Delivery in Women with Secondary Tubal Infertility

Marsha E. Wolf, PhD, Janet R. Daling, PhD, and Lynda F. Voigt, PhD

Abstract: The history of cesarean delivery was evaluated in a population-based case-control study of secondary infertility in King County, Washington. Sixty-one married women diagnosed with secondary infertility due to tubal problems who had a previous viable pregnancy were compared to 343 married women who had a previous viable pregnancy and then had a live birth that was conceived at the same time the infertile women began trying to conceive. The risk of tubal infertility was not substantially elevated in women who had a previous cesarean delivery in the most recent viable pregnancy compared to women with vaginal delivery (odds ratio = 1.2; 95% confidence interval = 0.4, 3.7). (*Am J Public Health* 1990; 80:1382–1383.)

Introduction

Cesarean section deliveries in the United States have increased from 5.5 per 100 deliveries in 1970 to 20.3 in 1983.[1,2] Women who have cesarean deliveries are at greater risk for postpartum complications of pelvic infections than women who deliver vaginally,[3] but the effect of cesarean delivery and related postpartum infection on subsequent fertility is not well understood.[4–10]

Methods

A case-control study design was used, in which the cases were 61 20–39 year old married women residents of King County, Washington who had secondary infertility (women who had previously conceived, but who were not able to conceive after trying for at least one year) and who have previously delivered a live birth or still birth. These women had their first medical evaluation for infertility during 1979 to 1981. (The study subjects were part of a larger study on infertility and contraception in which 89.5 percent of all physicians in King County who perform infertility evaluations agreed to participate.)[11] If cesarean section via a possible infection did lead to infertility in some women, it would more likely be women whose infertility was related to a tubal condition.[12,13] For this reason only women with blocked tubes, diagnosed by hysterosalpingogram or surgery, were included as cases. Women with an abnormal hysterosalpingogram who were later found by surgery to have an intrauterine abnormality, rather than a tubal condition, were excluded.

The control group (n = 343), identified from state vital records, was composed of 20–39 year old married women residents of King County who had a previous birth and then subsequently delivered a newborn who was conceived during the same years in which the infertile cases started trying to become pregnant.

By structured interview, data for each study subject were collected on demographic characteristics, use of "social" drugs, tobacco, alcohol, personal hygiene, and medical, reproductive, contraceptive, and sexual histories.

Results

Although cases and controls were somewhat similar with regard to education, income, and age at reference date (the time at which both the infertile case and the control began to attempt to conceive), the cases had a larger percentage of racial minorities and were more likely to have been employed, to be a smoker, and to have had more sexual partners, a history of Dalkon Shield use, and their last viable pregnancy before 1975 than did the controls (Table 1).

A cesarean section delivery of the most recent birth was found in four cases (6.6 percent) and in 29 controls (8.3 percent). Using unconditional logistic regression to obtain maximum likelihood estimates of relative risk, we adjusted for the effects of age (less than 30, 30 or older) and year of last viable pregnancy (before 1975, 1975 or after). The adjusted odds ratio, an estimate of the relative risk, was 1.2 (95 percent confidence interval = 0.4, 3.7). Since the number of exposed cases was small, we also performed exact stratified analysis with EGRET statistical software which utilizes the method described by Mehta, Patel and Gray.[14] The exact odds ratio, adjusted for the same variables as above had slightly wider confidence intervals (OR = 1.2; 95% CI = 0.3, 3.8). Variables that were assessed but did not appreciably change the cesarean delivery risk estimates included race, employment type, smoking status, parity, education, income, age at first coitus, number of lifetime sexual partners, history of Dalkon Shield and other intrauterine device use, history of gonorrhea, pelvic inflammatory disease, douching, marijuana use, and census tract of residence.

In our study, no case or control with a cesarean delivery reported postpartum pelvic infection.

Discussion

Although this study revealed no association between cesarean delivery and subsequent secondary tubal infertility, the small sample size limited the study's ability to detect a possible modest increase in risk associated with cesarean delivery. The 95% CI suggests a relative risk of 3.7 or more is highly unlikely.

Another limitation of the data is that women with a prior successful pregnancy who had a cesarean delivery may be less motivated to seek help or to undergo the expense or discomfort of invasive techniques, i.e. hysterosalpingogram or laparoscopic surgery necessary to diagnose tubal infertility. To the extent this is so, ascertainment of cases of women with secondary tubal infertility among women with cesarean section will be incomplete, and the observed risk will underestimate the true risk.

Finally, we were unable to assess the more direct measure of post-cesarean pelvic infection itself, except by self-report. Medical charts were not reviewed.

From the Department of Epidemiology, School of Public Health and Community Medicine, University of Washington, Seattle (all authors); the Harborview Injury Prevention and Research Center (Wolf); and the Fred Hutchinson Cancer Research Center, Division of Public Health Sciences (Daling, Voigt). Address reprint requests to Marsha E. Wolf, PhD, Harborview Injury Prevention and Research Center, 325 Ninth Avenue, Mailstop ZX-10, Seattle, WA 98104. This paper, submitted to the Journal October 24, 1988, was revised and accepted for publication March 20, 1990.

Table 1—Characteristics of Women with Secondary Tubal Infertility and Controls

Characteristic (at reference date)	Cases		Controls	
	No.	(%)	No.	(%)
Education				
<College	24	(39.3)	133	(38.8)
College	37	(60.7)	210	(61.2)
Annual Income[a]				
<$15,000	10	(16.9)	56	(16.4)
$15,000–30,000	30	(50.8)	180	(52.8)
>$30,000	19	(32.2)	105	(30.8)
Age Distribution (years)				
<20	3	(4.9)	1	(0.3)
20–24	9	(14.8)	56	(16.3)
25–29	32	(52.5)	159	(46.4)
30–34	14	(23.0)	110	(32.1)
35+	3	(4.9)	17	(5.0)
Race				
White	48	(78.7)	310	(90.4)
Black	7	(11.5)	10	(2.9)
Other	6	(9.8)	23	(6.7)
Employment Type				
None	16	(26.2)	177	(51.6)
Executive	3	(4.9)	9	(2.6)
White Collar	27	(44.3)	128	(37.3)
Blue Collar	15	(24.6)	29	(8.5)
Smoking Status				
Current	26	(42.6)	88	(25.7)
Never	28	(45.9)	192	(56.0)
Prior	7	(11.5)	63	(18.4)
Total Number of Lifetime Sexual Partners[b]				
1–4	39	(63.9)	263	(76.9)
≥5	22	(36.1)	79	(23.1)
Prior Dalkon Shield Use				
Yes	12	(19.7)	30	(8.7)
No	49	(80.3)	313	(91.3)
Reference Year				
<1976	8	(13.1)	21	(6.1)
1976–1978	25	(41.0)	128	(37.3)
≥1979	28	(45.9)	194	(56.6)
Parity				
1	50	(82.0)	266	(77.6)
2	7	(11.5)	55	(16.0)
3	4	(6.5)	14	(4.1)
≥4	0		8	(2.4)
Year of last viable pregnancy				
Before 1975	38	(62.3)	91	(26.5)
1975 or after	23	(37.7)	252	(73.5)

[a] Two cases and two controls did not answer this question.
[b] One control refused to answer.

Despite these limitations, this study adds to our knowledge; it is the first reported population-based case-control study which assesses the relative risk of secondary infertility following cesarean section. Our findings corroborate those of Hemminki, *et al*,[4] and Zdeb, *et al*,[6] who found essentially no difference in the proportion of women who had at least one subsequent child after the first pregnancy when comparing women with previous cesarean and vaginal deliveries.

The findings are in contrast to those found by La Sala[7] and Hemminki.[5] Of all the reported studies, only La Sala adequately excluded women who had been sterilized, were using contraceptives, or were not attempting a pregnancy. However, La Sala defined secondary infertility as two years of infertility, a definition not widely used in the United States.

Three other studies that addressed this topic found similar rates of subsequent pregnancies in women with emergency cesareans[9] or cesareans with subsequent infection[8,10] compared to other women with elective cesarean or cesarean without infection respectively.

One explanation for why a large increase in risk of infertility was not found in our study may be prophylaxis or early diagnosis and treatment of postpartum infections in women with cesarean deliveries which could prevent the severity and course of infection and thus decrease the occurrence of tubal damage.

It will take larger studies or several studies to be confident in these results. It is nonetheless reassuring that in a metropolitan area of nearly 1.3 million people in which most of the incident cases of infertility who sought help were identified, only four of the 61 cases of secondary tubal infertility with a prior live birth had a cesarean section in the last delivery, a rate similar to the comparison group and to obstetric patients seen during the same time period.

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A Comparison of Homeless, Community-wide, and Selected Distressed Samples on the CES-Depression Scale

Ferris J. Ritchey, PhD, Mark La Gory, PhD, Kevin M. Fitzpatrick, PhD, and Jeffrey Mullis, MA

Abstract: Center for Epidemiological Studies (CES) Depression Scale results for surveys of homeless, community-wide and selected distressed samples are compared. Nearly four times the percentage of homeless fit the criterion for clinical caseness (a score of 16+) compared to the general population (74 to 19 percent). None of the distressed samples exhibited a higher rate except psychiatric patients diagnosed as acutely depressive. High rates of depression have implications for social policies directed toward homelessness. (*Am J Public Health* 1990;80:1384–1386.)

Introduction

Psychological depression is characterized by depressed mood, negative self-concept, disturbed vegetative functioning, agitation, slower activity levels, distractibility, and indecisiveness.[1] These symptoms of extreme distress are especially detrimental for problem solving and coping among homeless persons with few financial and social resources, and limited control over their environments.

The Center for Epidemiological Studies Depression Scale (CES-D)[1-3] is a survey instrument scored from 0 to 60 (20 symptoms coded 0 to 3 depending on their frequency during the week prior to interview). It has high internal consistency,[1] strong test-retest reliability[4] and, although designed to measure depressive symptomatology in community populations, it is a sensitive predictor of diagnosed depression.[4-7] For example, Boyd, *et al*, found 80 percent correspondence between diagnosed depression and "possible clinical caseness" (a CES-D score of 16+).[8] The scale's specificity is strong; over 91 percent were correctly predicted to have nondepressive diagnosis (using the Diagnostic Interview Schedule) when the 21+ cutoff was used for the CES-D.[4] For our homeless sample, the Cronbach's alpha reliability coefficient for the CES-D was .89; item-to-total correlations ranged from .24 to .74.

Using the CES-D Scale,[1-3] we compared rates of depressive symptoms for a sample of homeless persons to existing data on community-wide and selected distressed samples.

Methods

Data on the homeless were from the Birmingham (Alabama) Homeless Enumeration and Survey Project conducted in early 1987.[9] The Metropolitan Statistical Area (MSA) population was approximately 900,000. The enumeration phase was a single-night census of homeless persons finding 498 in MSA shelters, and 103 on the streets downtown where the homeless were known to be concentrated. An interview survey of 150 adults followed in the three months after the enumeration. A random probability sampling design was used with quotas for the variables sex, race, and geographic site (i.e., shelter versus street) based on enumeration data. Whites were slightly overrepresented in the street sample. Interviews averaged 62 minutes. Of 155 people asked for interviews, only one refused. Four interviews were unusable.

Results

Table 1 presents percentages of possible clinical caseness (CES-D of 16+) for community-wide, selected distressed, and homeless samples.* Across five general population studies in various communities between 1971 and 1985, the results fell consistently below 20 percent. In the distressed samples (discounting acute depressives), percentages ranged from 23.0 (physically or psychologically impaired persons over 55) to 79.0 (rural mental health clinic patients). Generally, the distressed samples comprised of psychiatric patients had higher rates than samples characterized by stressful life circumstances.

Comparing community-wide and distressed samples to our sample revealed a considerable amount of distress among the homeless. Only one respondent of our sample was asymptomatic. Nearly three of four (73.3 percent) scored 16 or above. This percentage was similar to Koegel and Burnam's[14] finding of 71 percent clinical caseness among homeless males of inner-city Los Angeles. In addition, we found 37 percent to score 30 or above, a result similar to Susser's 33 percent for a sample of male first-time shelter users primarily under 40 years of age.[19] Except for Weissman's study of clinic patients diagnosed with acute depression,[6] even groups as distressed as psychiatric inpatients, alcoholics under treatment, and rural mental health clinic patients did not have significantly higher rates.

Comparisons to domiciled samples with stressful life circumstances were informative. For example, low income mothers of young children had significantly lower rates than the Birmingham homeless, as did the unemployed, suggesting that environmental and situational stressors of homelessness exact a greater toll on mental well-being than those faced by the housed poor. Finally, for persons recently widowed or separated—two highly stressful life events—the rates of clinical caseness were lower than that of the homeless sample. An analysis elsewhere of variation in depressive symptomatology within our homeless sample revealed that homelessness does not impact its victims differentially.[20] No matter how the sample was broken down, the mean CES-D score for any subsample was more than twice the mean score of any of the community-wide samples.**

From the Department of Sociology, University of Alabama at Birmingham. Address reprint requests to Ferris J. Ritchey, PhD, Associate Professor, Department of Sociology, School of Social and Behavioral Sciences, University of Alabama, UAB Station, Birmingham, AL 35294. This paper, submitted to the Journal November 9, 1989, was revised and accepted for publication March 26, 1990.

*Analysis of mean differences in CES-D scores yielded the same results. These data are available upon request from the authors.

**A table summarizing mean differences of sample subgroups is available upon request from the authors.

TABLE 1—Comparison of CES-D Scores: Community-Wide, Selected Distressed, and Homeless Samples

Studies and Year(s)	Research Site/Sample/Subsamples	N	Percent CES-D 16 or above
Community-Wide Samples			
Radloff;[1] Comstock & Helsing[5] (1971–73)	Washington County, Maryland & Kansas City, Missouri	2845	19.5^{**}
Eaton & Kessler;[10] Sayetta & Johnson[11] (1971–75)	National Health & Nutrition Examination Survey	2867	16.4^{**}
Frerichs, et al.[12] (1979)	Multi-Ethnic sample, Los Angeles	1000	19.1^{**}
Ensel;[4] Lin and Ensel[13] (1979–80)	Tri-County Area, Upstate New York	871	16.9^{**}
Koegel and Burnham[14] (1984–85)	Inner-City Los Angeles, Males only	1466	9.4^{**}
Selected Distressed Samples			
Radloff[15] (1971–73)	Washington County, Maryland & Kansas City, Missouri		
	Psychiatric Inpatients	70	70.0
	Death of spouse-past year	33	51.5^{*}
	Marital separation-past year	63	49.2^{**}
	Unemployed	199	30.2^{**}
Weissman, et al.[6] (1977)	Connecticut Mental Health Center, New Haven		
	Acute Depressives	148	99.3^{**}
	Recovered Depressives	87	43.7^{**}
	Drug Addicts	60	51.7^{**}
	Alcoholics	61	62.3
	Schizophrenics	50	36.0^{**}
Hall, et al.[16] (1976–77)	Greensboro, North Carolina; Low-Income Mothers of Young Children	111	48.0^{**}
Husaini, et al.[17] (1977)	9 Rural Counties, Tenn.; Mental Clinics and Veterans Administration Hospital	194	79.0
Orr, et al.[7] (1984)	Baltimore; Mothers of Young Children	149	38.3^{**}
Frerichs, et al.[12] (1979)	Multi-Ethnic sample, Los Angeles		
	Unemployed	98	46.9^{**}
	Low Total Family Income ($0–$8,499)	239	29.3^{**}
Another authors' sample[a] (1986–87)	4 Metropolitan Counties in Alabama: Physically and/or psychologically impaired adults, 55 years and older	154	23.0^{**}
Homeless Samples			
Koegel and Burnam[14] (1984–85)	Inner-City Los Angeles, Males only	379	71.0
Rossi, et al.[18] (1987)	Chicago	319	47.0[b]
Present study (1987)	Birmingham, Alabama	150	73.3

[a] FROM: La Gory M, Fitzpatrick KM: The effects of environmental context on elderly depression. Presented to the American Gerontological Society Meetings, Minneapolis, Minnesota, 1989.

[b] These figures cannot be compared directly with others. Only six CES-D items were used.

*Significantly different from Birmingham homeless, $p < .05$, two-tail test.

**Significantly different from Birmingham homeless, $p < .001$, two-tail test.

Discussion

The homeless suffer from high rates of depressive symptomatology indicative of extreme distress. Only acute depressives, mental health clinic patients, psychiatric inpatients, and alcoholics have rates of clinical caseness on the CES-D comparable to the homeless. Some caveats are in order. First, given the extreme environmental conditions of homelessness (e.g., poor nutrition, climatic exposure), the CES-D may be tapping physiological as well as psychological dimensions.[6] Second, the CES-D may be, for some respondents, gauging reactions to acute situational factors rather than gauging depressive illness or chronic demoralization.[13,21-24] Third, homelessness is usually accompanied by other stressful life events and circumstances; thus, depression may precede homelessness in some cases.[24] Notwithstanding these considerations, depressive symptoms have implications for an individual's ability to resolve the situation, even if external conditions (e.g., available jobs and housing) are favorable. These results suggest that current shelter, rehousing, and treatment programs should institute intervention strategies to avert the potentially cyclical nature of homelessness and depression.

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How Valid Are Mammography Self-Reports?

EUNICE S. KING, RN, PHD, BARBARA K. RIMER, MPH, DRPH, BRUCE TROCK, PHD,
ANDREW BALSHEM, BA, AND PAUL ENGSTROM, MD

Abstract: We compared mammography reports in medical records to self-reports obtained during a 1989 telephone interview survey for a sample of 100 women members of a health maintenance organization (HMO) who indicated they had mammograms within the past year and 100 who said they had not had mammograms within the past year. Of the women reporting they had not had mammograms within the past year, none had mammogram reports in the HMO data center. Of the 100 women reporting they had mammograms within the past year, 94 had confirmatory radiology records. (*Am J Public Health* 1990; 80:1386–1388.)

Introduction

Programs of regular mammography screening can produce downward shifts in the stage of breast cancer and reduce mortality by as much as 35 percent.[1–4] The National Cancer Institute (NCI), the American Cancer Society, and other major medical organizations recommend annual screening mammograms for all women ages 50 and older and screening every one to two years for women ages 40–49 years.[5] A potential problem in evaluating the effectiveness of programs aimed at increasing mammography utilization is that most surveys and program evaluations have relied exclusively upon women's self-reports of mammograms. The draft of the US Department of Health and Human Services' Year 2000 Objectives for the Nation stipulates that one of the key assumptions underlying the tracking of changes in mammography utilization is that self-report data provide a valid measure of screening utilization.[6]

If mammography self-reports are not accurate, survey data may overestimate or underestimate actual mammography utilization. Although the accuracy of self-report data in such diverse areas as cervical cancer screening and smoking cessation has been studied,[7–11] the findings cannot be generalized from one content area to another.[7] Based on a MEDLINE search, we found no published studies in which mammography self-reports were validated. Therefore, we undertook the present study.

Methods

This validation study was part of the Avoidable Mortality Study,[4] conducted in conjunction with US Healthcare Check, a free breast screening program offered to all women members of an IPA-model (independent practice association) HMO (health maintenance organization) ages 40 years and over. National guidelines are followed: mammograms are recommended every one to two years for women ages 40 to 49 years and annually for women age 50 and older.[5]

Four telephone survey waves are conducted as part of the Avoidable Mortality Study to evaluate the effectiveness of the program in increasing mammography utilization among women in the HMO compared to geographic controls. US Healthcare Check maintains a centralized mammogram report database. Radiology reports of HMO members are sent to the US Healthcare Check office and to the women's physicians.

In January 1989, a telephone survey was conducted as part of the Avoidable Mortality Study. A sample of 450 HMO women between the ages of 50 and 74 was selected randomly from approximately 45,000 eligible women to whom the US Healthcare Check program is targeted. Interviews were completed with 84 percent; fewer than 10 percent of the women refused to be interviewed.

During the 15-minute structured telephone interview conducted by professional interviewers, women who said they had mammograms were asked when the procedure was most recently done. Mammography utilization questions were those used in the National Health Interview Survey (NHIS) 1987 Cancer Control Supplement and by members of the NCI's Breast Screening Consortium.[4] Overall, 63 percent of the HMO respondents said they had a mammogram in the past year. Mammography rates similar to those in our study and higher than the NHIS[12] also were found in surveys

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From the Fox Chase Cancer Center. Address reprint requests to Eunice S. King, RN, PhD, Population Science, Fox Chase Cancer Center, 510 Township Line Road, Cheltenham, PA 19012. This paper, submitted to the Journal October 16, 1989, was revised and accepted for publication April 10, 1990.

conducted by five other NCI Mammography Consortium sites.[13] We questioned whether overreporting could account for some of these differences.

In August 1989, we randomly selected 100 women from the 278 HMO women whose responses indicated they *had* a mammogram within the past year and 100 HMO women from the 171 who indicated they *had not had* a mammogram within the past year. These women comprised the validation study sample.

For women reporting having had mammograms within the past year, project staff contacted the HMO assigned radiology center if the database contained no confirmatory record. If the center was unable to verify the woman's mammogram through radiology reports, the primary physician was identified through the database and contacted. If the physician had no record of the mammogram, the woman herself was contacted to provide additional information about where and when she had the mammogram. Subsequently, with the woman's consent, study staff contacted the radiology center indicated for confirmation of the mammogram. In all instances, women's self-reports were considered valid only if they reported having had a mammogram, and this report was validated by database information or confirmed by radiology or physician's office records.

Results

Of the 200 women originally selected for the validation study, one had to be eliminated due to a data entry error. In general, women reporting having had a mammogram within the past 12 months did not differ from those reporting not having had a mammogram within the past 12 months. However, working women were significantly less likely to have obtained mammograms, and there were more employed women in the group who had not had mammograms (Table 1).

Women's self-reports of having obtained mammograms within the past year correlated highly with the presence of actual mammogram reports (Table 2). Of the 99 women reporting they had *not* had mammograms within the past year, none had a mammogram report in the database. Of the 100 women reporting they "had a mammogram" within the past year, we were able to validate self-reports on 94 women. Mammogram reports on 50 of these women were obtained through the database; the remaining 44 were validated through phone calls to radiology centers or to the women's physicians. Although *all* of the remaining six women had obtained mammograms, they were not done within the past year. Five of the women had mammograms between 13 and 18 months prior to the survey interview; the other woman had a mammogram two years prior to the survey. These six women did not share any particular sociodemographic characteristics when examined by age, employment, race, education, and marital status. Thus, ultimately, we were able to account for *every* one of the women who reported that they had mammograms.

Discussion

These data indicate agreement between women's self-reports of mammography and providers' records. But, for the 99 women reporting they had *not* had mammograms within the past year, we did not attempt further verification beyond the data in the US Healthcare Check database, as we did for the women reporting they had obtained mammograms within the past year.

TABLE 1—Demographic Characteristics of Women in the Mammogram Validation Study

	Mammogram in last 12 months				
	Yes		No		
Characteristics	n	(%)	n	(%)	*p*
Age (years)					
50–54	24	(24)	29	(29)	
55–59	30	(30)	32	(33)	
60–64	16	(16)	20	(20)	
65–69	19	(19)	11	(11)	
70–74	11	(11)	7	(7)	
Total	100	(100)	99	(100)	$p = .40$
Race					
White	80	(80)	80	(81)	
Non-White	20	(20)	19	(19)	
Total	100	(100)	99	(100)	$p = .80$
Annual Income					
<$15,000	22	(28)	19	(22)	
$15,000–$30,000	31	(39)	32	(38)	
>$30,000	20	(33)	34	(40)	
Total	79*	(100)	85*	(100)	$p = .58$
Marital Status					
Married	72	(73)	62	(64)	
Not Married	26	(27)	36	(37)	
Total	98*	(100)	98*	(100)	$p = .14$
Education					
<High School Graduate	21	(21)	25	(25)	
High School Graduate	50	(51)	45	(46)	
>High School Graduate	28	(28)	28	(28)	
Total	99*	(100)	98*	(100)	$p = .77$
Employment					
Employed	57	(58)	74	(76)	
Not employed	42	(42)	24	(24)	
Total	99*	(100)	98*	(100)	$p = .01$

*Numbers do not equal total n due to refusals or missing data.

It is not feasible to verify such negative reports beyond the HMO database, since it would require contacting every physician who could have ordered a mammogram for each woman. At first, only 50 percent of women who reported having a mammogram had verification in the HMO database. These omissions were due to the radiologist failing to report the procedure in a timely fashion; the women had responded correctly. For a similar rate of omission to exist among the women reporting *not* having mammograms would also require an error in self-report on the part of the majority of women. Given the high rate of accuracy among women reporting mammograms, a high rate of inaccurate response among women reporting no mammogram seems highly unlikely.

In a study by Sawyer, *et al*,[11] of the accuracy of women's self-reports of Pap tests, the underreporting errors—although less than the overreporting errors of 21 percent—were 17

TABLE 2—Comparison of Women's Self-reports with Actual Mammogram Reports

	Radiology Reports		
Self-reports	Yes n	No n	Total n
Had mammogram in preceding year	94	6	100
Did not have mammogram in preceding year	0	99	99
Total	94	105	199

Phi coefficient = .94 (95% CI = 0.93, 9.95), $p < .001$.

percent. Unlike having a Pap test, which is usually taken only during a pelvic examination, and sometimes without the woman's knowledge, having a mammogram is a discrete event which occurs at a time or a place other than the medical examination and thus is less likely to be confused or forgotten.

Although further studies in other populations should be done to confirm the results found here, these data provide a basis for optimism. If the findings are replicated, it appears that women's self-reports can be used accurately to monitor national changes in mammography utilization.

ACKNOWLEDGMENTS

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Incidence of Bicycle-Related Injuries in a Defined Population

DIANE C. THOMPSON, MS, ROBERT S. THOMPSON, MD, AND FREDERICK P. RIVARA, MD, MPH

Abstract: Population-based incidence rates for head injuries and total injuries resulting from bicycle crashes were calculated in a Seattle, Washington health maintenance organization population. Overall rates were 163 per 100,00 for all injuries and 42/100,000 for head injuries. Individuals between 5 and 14 years of age are at highest risk for bicycle-related injuries. The data are presented for their potential utility in program planning. (*Am J Public Health* 1990; 80:1388–1389.)

Introduction

Bicycling is an increasingly popular sport in the United States. Injuries due to bicycling accounted for approximately 574,000 emergency room visits and 1,300 deaths in 1985, with head injuries accounting for the major morbidity and mortality.[1,2] Helmets have been shown to be an effective means of preventing these injuries.[3] Programs to promote helmet use have been successful.[4] The present study was undertaken to provide information on the incidence of these injuries for use in community bicycle helmet campaigns and safety programs.

Methods

Injured cyclists were identified during a one year surveillance of hospital emergency rooms in Seattle, Washington (December 1, 1986 to November 30, 1987). The methods are detailed elsewhere.[3] Cyclists for this study are limited to those treated at the facilities serving Group Health Cooperative of Puget Sound (GHC), a large staff-model health maintenance organization, thus allowing the calculation of population-based rates. GHC membership is demographically similar to the surrounding population in the Seattle metropolitan area, but differs from the US population by its greater educational level and underrepresentation of Blacks. Group Health Cooperative members are highly educated; 67 percent have more than a high school education. The ethnic composition of GHC membership is 91 percent Caucasian, 3 percent Black, 4 percent Asian/Pacific Islander, and 2 percent other, compared to 83.4 percent, 11.7 percent, 1.7 percent, and 3.2 percent, respectively in the US population.

GHC Central and East regions include 223,298 members who receive nearly all their outpatient, emergency, and hospital care at GHC facilities. There were two severely injured cyclists treated at the regional trauma center; these were included in the study. Review of the medical examiners' records for the study period indicated no deaths at the

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Address reprint requests to Diane C. Thompson, MS, Research Coordinator, Harborview Injury Prevention and Research Center, 3225 Ninth Avenue, ZX-10, Seattle, WA 98104. Dr. R. Thompson is Senior Director of Preventive Care Research, Department of Preventive Care, Group Health Cooperative of Puget Sound; Dr. Rivara is Director of Harborview Injury Prevention and Research Center, as well as Professor of Pediatrics and Adjunct Professor of Epidemiology, University of Washington School of Medicine. This paper, submitted to the Journal December 11, 1989, was revised and accepted for publication April 9, 1990.

accident scene. Bicyclists injured outside the Seattle metropolitan area or treated at other area medical facilities were not included.

Exposure to cycling was measured by asking the frequency of riding and the number of miles and hours cycled per week. Self-reported mileage figures were used to calculate the mean number of miles cycled per week, and to estimate injuries per 100 miles ridden.

Incidence rates of emergency room treated injuries resulting from bicycle crashes were calculated using the mid-year 1986 population of the GHC East and Central regions.

Results

There were 364 Group Health members treated in the emergency room for bicycle-related injuries. One hundred-one (28 percent) were head injured. Thirty (29.7 percent) of the head injured experienced brain injuries; 4 percent were hospitalized. The 364 cyclists constituted 54.5 percent of all cyclists seen in the five area hospitals.

Table 1 displays age- and sex-specific injury rates. Head injury rates were highest in the 5–9 year age group while rates for all bicycle injuries were highest in the 10–14 year age group. Males had higher injury rates than females. When the incidence of bicycle-related injuries was age-adjusted to the US population, the rate was 51 per 100,000 for head injuries and 187 per 100,000 for all bicycle injuries.

Fifty-six of the 364 injured cyclists (15.4 percent) were wearing helmets at the time of the crash; only seven of the 56 head-injured were helmeted. Helmet use was very low among riders under 15 years of age (4.3 percent), but increased to 23 percent among 15–24 year olds and to 44 percent for those 25 years of age and older. There were no significant differences in rates of helmet use by sex.

Exposure to cycling is summarized in Tables 2 and 3. The injury rate per 100 miles ridden per year is highest for 5–14 year olds. Seventy-seven percent of the ER visits were between April and September.

TABLE 1—Incidence of Bicycle-Related Head Injuries and All Injuries per 100,000 population by Sex

Age (years)	Injuries (N)	Sex		Total	SE*
		Male	Female		
Head Injuries					
0–4	10	94.3	41.9	68.6	22
5–9	43	414.7	147.1	283.0	43
10–14	27	295.1	72.6	188.3	36
15–19	5	59.9	—	30.2	14
20–24	3	38.6	14.9	25.2	15
25–44	10	14.8	21.1	13.3	4
45–64	3	14.3	—	6.7	4
All ages	101			42.2	
All Bicycle Injuries					
0–4	12	121.2	41.9	82.3	24
5–9	102	907.1	414.6	671.3	66
10–14	116	1260.9	246.9	809.0	75
15–19	43	443.6	72.9	259.6	40
20–24	18	250.6	74.5	151.3	36
25–44	55	109.7	43.4	73.2	10
45–64	17	52.3	25.3	38.0	9
All Ages	364			163.0	

*Standard Error

TABLE 2—Mean Miles per Week Cycled and Injuries per 100 Miles Ridden per year According to Age and Type of Injury

Age Group (years)	Mean Miles per Week Cycled			Injuries per 100 Miles Ridden per Year
	Head Injuries	Other Injuries	All Injuries	
0–4	5.7	3.5	5.3	4.4
5–9	7.4	6.0	6.6	30.0
10–14	10.3	12.8	12.3	18.1
15–19	10.0	48.7	43.8	1.9
20–24	71.3	40.9	45.9	0.8
25–44	47.8	65.7	63.0	1.7
45–64	9.0	56.9	47.9	0.7
65+	—	—	1.0	—

TABLE 3—Distribution of Season and Week of All Bicycle Injuries

Season Injured	%	Day of Week Injured	%
January–March	13.0	Saturday/Sunday	41.0
April–June	42.0	Monday	10.2
July–September	35.3	Tuesday	11.3
October–December	9.1	Wednesday	9.1
		Thursday	16.6
		Friday	11.6

Discussion

Although helmet use has a direct effect on head injuries,[3] it does not explain why the peak ages of head injuries and overall injuries are different or why the ratio of head to all injuries is higher for young children. Our cycling exposure data, which are similar to those of other studies[1,5–7], do not explain the differences. Possible explanations are differential ascertainment of injury types and increased injury severity in older children.[8]

We assume most cycling mileage was estimated based on distance between destinations rather than actual measurement. The number of bicycle odometers used is unknown. Mileage data indicate children 5–9 years old are at highest injury risk.

Friede and co-workers (SCIPP) reported 87.8 bicycle injuries per 10,000 person years (0–19), with the injury rate highest in 6–12 year old males.[5] Our overall injury rate for this age group was 45 per 10,000. Reasons for the discrepancy are not clear. Lack of case ascertainment in the present work does not appear to be the explanation because it was a closed system of care. Other potential explanations for the differences include SCIPP'S use of all pedal cycle accidents (bicycles, tricycles, other pedal powered vehicles), sampling methods (SCIPP used a 25 percent sample of accidents over a two-year period),[5] and the proportion of all types of GHC injuries treated in the ER (40 percent) versus primary care (60 percent).[8] This latter study indicated that GHC overall rates (clinic and ER combined) appear closer to ER rates obtained by SCIPP[9] and NEOTS.[10] Two other studies[11,12] reported incidence rates of medically treated injuries similar to rates at GHC.[8] Waller (1971) reported higher bicycle injury rates. His denominator, all bicycle owners, yields higher rates than using the general population.[7]

Brain injury rates are comparable to other studies. Kraus, *et al*, reported population-based rates of brain injuries from bicycle crashes requiring hospitalization of 13.5 per 100,000 with the highest injury rates in children under age 15.[13] The rate for all brain injuries from the present work was 12.5 per 100,000.

The rates of injuries reported here are conservative since they represent only ER treated injuries. Many injured cyclists, possibly the majority, were treated at home or in outpatient clinics. Bicycling injuries contribute substantially to health care utilization and are a chief cause of recreational injury.[1,6,14–20] Children age 5–14 are at greatest risk of injury, and are an appropriate target for the application of available and effective prevention programs.[2,3,5,21]

The data presented are offered for their potential utility to the communities planning bicycle safety helmet campaigns.

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Different Views

Hypertension Prevalence and the Status of Awareness, Treatment, and Control in the Hispanic Health and Nutrition Examination Survey (HHANES), 1982–84

Gregory Pappas, MD, PhD, Peter J. Gergen, MD, MPH, and Margaret Carroll, MSPH

Abstract: The prevalence rates of hypertension among adult (ages 18–74) Mexican Americans, Cuban Americans, and Puerto Ricans were estimated using data from the 1982–84 Hispanic Health and Nutrition Examination Survey (HHANES). Hypertension is defined as diastolic greater than or equal to 90 mm Hg, or systolic greater than or equal to 140 mm Hg, or currently taking antihypertensive medication. Among Mexican Americans in the Southwestern United States, 16.8 percent of the males and 14.1 percent of the females were found to be hypertensive. Among Cuban Americans in Dade County, Florida 22.8 percent of the males and 15.5 percent of the females were hypertensive. Among Puerto Ricans in the New York City area 15.6 percent of the males and 11.5 percent of the females were hypertensive. The age-adjusted rates are significantly lower than comparable rates for Whites and Blacks as measured in the second National Health and Nutrition Examination Survey (NHANES II), 1976–80. Control of hypertension in the HHANES populations fall short of the 1990 Objectives for the Nation established by the US Public Health Service 60 percent (34 percent controlled Mexican American hypertensives, 27.8 percent controlled Cuban American hypertensives, and 29 percent controlled Puerto Rican hypertensives). (*Am J Public Health* 1990; 80:1431–1436.)

Introduction

Hypertension continues to be a leading contributor to morbidity and mortality in the United States.[1] While the number of studies on the extent and nature of hypertension in the US population has increased, our knowledge of how this condition affects Hispanics has lagged behind.[2] Hispanics are the second largest minority in the United States and constitute a rapidly growing population.[3] An assessment of risk factors for hypertension has suggested that Hispanics may have high rates of the condition.[2] Surveys which have estimated the prevalence of hypertension among Hispanics have given inconclusive results.[4–8]

The drive toward better blood pressure control has been heightened by the success measured in a decline in stroke and coronary artery disease mortality.[1] The US Public Health Service has set "1990 Objectives for the Nation" including nine dealing with hypertension control. The 1990 Objectives directed towards improving health status of hypertensives state that ". . . at least sixty percent of the estimated population having definite high blood pressure (160/95) or higher should have attained successful long-term blood pressure control, i.e. a blood pressure at or below 140/90 for two or more years."[9] The extent to which Hispanics have met these goals is unknown.

This paper provides estimates of the prevalence of hypertension among three Hispanics subgroups based on the Hispanic Health and Nutrition Examination Survey (HHANES) conducted by the National Center for Health Statistics between 1982 and 1984. These estimates are compared to the published results for Whites and Blacks in the general population using the second National Health and Nutrition Examination Survey (NHANES II), 1976–80.[10] The status of awareness, treatment, and control of hypertension among Hispanics has also been estimated using HHANES data.

Methods

Survey Design

The HHANES was conducted among three Hispanic subgroups in the United States from July 1982 through December 1984. The three target populations consisted of the civilian, noninstitutionalized population of Mexican origin or ancestry in selected parts of five southwestern states (California, Arizona, New Mexico, Colorado, Texas); of Cuban origin or ancestry in Dade County (Miami), Florida; and of Puerto Rican origin or ancestry in selected counties in New York, New Jersey, and Connecticut.

The survey design of HHANES was a stratified, multistage, probability sample of persons ages six months through 74 years. Selected households within the survey were visited by a bilingual interviewer who identified eligible Hispanic families. A pre-determined sampling pattern was followed to choose specified family members to participate in the survey. A family was considered eligible if the national origin or ancestry of at least one family member met the criteria (Mexican American in the southwest, Cuban American in

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From the National Center for Health Statistics, Dept. of Health & Human Services, US Public Health Service. Address reprint requests to Gregory Pappas, MD, PhD, National Center for Health Statistics, 6525 Belcrest Rd., Hyattsville, MD 20782. This paper, submitted to the Journal May 22, 1989, was revised and accepted for publication January 2, 1990.

Editor's Note: See also related Different View p 1437 and editorial p 1427.

Acronyms Used:

DBP:	Diastolic blood pressure
SBP:	Systolic blood pressure
HHANES:	Hispanic Health and Nutrition Examination Survey
JNC 1984:	Joint National Committee on Detection, Evaluation and Treatment of High Blood Pressure 1984
JNC 1988:	Joint National Committee on Detection, Evaluation and Treatment of High Blood Pressure 1988
NHANES II:	Second National Health and Nutrition Examination Survey

Dade County, and Puerto Rican in the New York City area); however, the selected sample persons themselves were not necessarily of the same origin. Analyses in this paper are based on data only for those sample persons of Mexican origin or ancestry in the southwest, of Cuban origin or ancestry in Miami, Florida, and of Puerto Rican origin or ancestry in the New York City area.

Details of the HHANES sample design, selection process, operational plan, quality control procedures, and questionnaires have been previously documented.[11]

Data Collection

The HHANES included both a household interview and a physical examination conducted in a specially designed mobile examination center. As part of the household interview, sample persons were asked if they had ever been told by a physician or other health professional that they had hypertension and if they were currently using medication for treating hypertension. Upon completion of the household interview, participants were scheduled for a physical examination that included blood pressure measurement, conducted in the mobile examination center.

Systolic (first phase) and diastolic (fifth phase) blood pressure were measured to the nearest even digit using a standard mercury sphygmomanometer. Two blood pressure measurements were taken on one occasion in the mobile examination center as part of a physician's examination. Both measurements were taken with the patient seated, five minutes into the examination and five minutes apart. The average of the two readings was used for the estimates presented here.

Definitions

For the purpose of the prevalence estimates presented in this paper, persons are considered hypertensive if they meet one of the following conditions set by the 1988 Report of the Joint National Commission on Detection, Evaluation, and Treatment of High Blood Pressure (JNC 1988): 1) a diastolic blood pressure of greater than or equal to 90 mm Hg; 2) a systolic blood pressure greater than or equal to 140 mm Hg; 3) reporting currently using anti-hypertensive medication (regardless of blood pressure measurements).[1] The 1990 Objectives use a second definition for hypertension setting the thresholds at diastolic greater than or equal to 95 mm Hg and systolic greater than or equal to 160 mm Hg. This second definition for hypertension (160/95) is used to assess the levels of awareness, treatment, and control of hypertension.

Prevalence estimates for the specific categories of adult blood pressure levels follow the JNC 1988 guidelines. These categories are listed and defined in the Appendix. Blood pressure levels are reported as a percent of the HHANES target population and as a percent of the hypertensives in those populations.

Hypertension prevalence estimates are available for Whites and Blacks in the general population based on the NHANES II (1976–80). These estimates have been previously published in a report of the Joint National Committee on Detection, Evaluation, and Treatment of High Blood Pressure, 1984 (JNC 1984).[10] The JNC 1984 estimates for hypertension were based on the average of three pressures (taken with the examinee sitting, lying, and again sitting). A caveat for the comparisons is that the JNC 1984 estimates included a small number of Hispanics. However, the removal of Hispanics from the sample does not produce estimates that show statistically significant differences from the estimates published in JNC 1984 report.

The percent of hypertensives under control on medication is defined as: the number of hypertensives who report taking medication and whose blood pressure is less than the threshold of 140/90 mm Hg (numerator) divided by the total of the hypertensives (denominator). Control is reported as a proportion of both definitions of hypertension, using the 140/90 and 160/95 thresholds.

The higher threshold for hypertension determines the denominator for the status variables (awareness, treatment, control). A person is considered aware of their hypertension if they report having ever been told by a physician or other health professional that they have high blood pressure or hypertension. Those who report they are currently taking antihypertensive medication are considered treated.

Response Rates

Among those selected for HHANES and eligible (of Hispanic origin), ages 18 to 74 years, there were 5,130 Mexican Americans, 1,560 Cuban Americans, and 1,942 Puerto Ricans. Of the selected persons, the proportion of persons interviewed was 77.4 percent, 72.6 percent, and 77.9 percent for Mexican Americans, Cuban Americans, and Puerto Ricans, respectively. Two blood pressure measurements were obtained on 82.3 percent, 75 percent, and 77.7 percent of the interviewed persons in the Mexican American, Cuban American, and Puerto Rican samples, respectively. No significant differences were noted with regard to age, sex, and antihypertensive medication status between persons interviewed and persons with two blood pressure measurements in any of these subgroups.

Systematic Bias in BP Measurements

The tendency toward blood pressure readings ending in zero was identified as a problem in HHANES (diastolic 23.8 percent, systolic 27.6 percent). Further analysis of this systematic bias was made around the critical values used to define hypertension: 90 mm Hg diastolic and 140 mm Hg systolic. Diastolic 86 mm Hg and 88 mm Hg were read more often than either 90 mm Hg or 92 mm Hg. Systolic 140 was read more often than 138. Thus it appears that the examiners tended to read below the level defining hypertension on the diastolic reading and above the level defining systolic hypertension. Therefore, the previously noted systematic bias may be thought to have an equivocal effect on the point estimates presented here.

Statistical Analysis

Sampling weights were used when calculating point estimates in order to account for individual selection probabilities and adjustments for non-response, non-coverage, and poststratification which resulted from the complex survey design used in HHANES. When the appropriate weights are used, the samples are representative of the Mexican origin population in the five southwestern states, the Cuban origin population in Dade County, Florida, and the Puerto Rican origin population in the New York City area.

The complex survey design used in the HHANES tended to increase the variance which would have been obtained through simple random sampling.[12] The complex survey design was accounted for by using an average design effect to adjust the variance calculated under the assumption of simple random sampling. As described by Kovar,[13] the design effect is the ratio of the variance of a statistic from a complex sample to the variance of the same statistic from a simple random sample. The value of the design effect is an indicator of the impact of the complex sample design on the variance.

The average design effect for hypertension was 1.0 for all three Hispanic subgroups. The complex design, therefore, has little effect on the variances and one could consider assuming simple random sampling for these analyses. A t-test was used to compare point estimates. All data analyses were done using programs accessible through SAS.[14]

The direct method was used for age adjustment. The reference population was the civilian noninstitutionalized population of the United States at the midpoint of the NHANES II, March 1, 1978.

Results

Prevalence of hypertension from HHANES is presented by subgroup, sex, and age in Table 1. These estimates represent approximately 775,000 hypertensive Mexican Americans in the southwest, 56,400 hypertensive Cubans in Dade County, Florida, and 78,000 hypertensive Puerto Ricans in the New York City area at the 1983 population levels.

Expected age and sex differences in hypertension were observed for the three Hispanic subgroups. The prevalence of hypertension was higher in older age groups for each of the three Hispanic subgroups. The rates of hypertension were age-adjusted, using the US population in 1978 as a reference (the mid-point of NHANES II). Men had significantly higher age-adjusted prevalence estimates for hypertension compared to women in the Mexican American (95% CI = 0.2, 6.2) and Cuban American (95% CI = 1.4, 12) subgroups. Age-specific comparisons between sexes suggest the same conclusion.

Table 2 presents the distribution of blood pressure levels regardless of medication status, among men and women in the Hispanic subgroups. The results are similar for all three Hispanic subgroups. The great majority of individuals in the HHANES populations have blood pressure measurements in the normal range. Mild diastolic elevation (DBP 90-104) ranges from 2.7 percent to 11.1 percent in the HHANES populations. Isolated systolic hypertension (SBP > = 160) was discovered in a very small portion of the population (less than 1.2 percent).

The distribution of blood pressure levels among hypertensives is presented in Table 3. By far the largest numbers of hypertensives are found in the mild hypertensive category (DBP 90-104), approximately 40 percent. Taken together, those with borderline isolated systolic hypertension (SBP 140-159) and isolated systolic hypertension levels (SBP > = 160) made up over one-third of all the Hispanic hypertensives. Less than 20 percent of all hypertensives (140/90) were controlled by medication.

A number of statistically significant results were observed comparing hypertension rates from HHANES with Whites and Blacks in the general population as measured in NHANES II (1976–80). Hypertension is more prevalent among Whites or Blacks than it is among Mexican Americans, Cuban Americans, or Puerto Ricans (see Table 4). Individual comparisons using the age-adjusted estimates revealed that each of the three Hispanic subgroups had statistically significant lower levels of hypertension than did either Blacks or Whites. Comparisons of age-specific rates suggest the same conclusion.

TABLE 1—Hypertension* Prevalence by Hispanic Subgroup, Sex and Age

	Males			Females		
Sex, Hispanic Origin, and Age (years)	Number of examined persons	Percent with hypertension	Standard error of the percent	Number of examined persons	Percent with hypertension	Standard error of the percent
Mexican Americans						
18–74	1553	16.8	1.2	1965	14.1	0.9
18–74 age adjusted		22.9	1.1		19.7	0.9
18–24	342	5.1	1.6	417	0.5	0.4
25–34	423	10.4	2	527	4.9	1.1
35–44	246	12.3	2.8	333	7.6	1.7
45–54	268	28.5	2.8	350	24.1	2.2
55–64	191	47.2	3.6	219	44.9	3.4
65–74	83	59.6	6.4	119	66.8	4.6
Cuban Americans						
18–74	398	22.8	1.2	498	15.5	0.9
18–74 age adjusted		20.5	2		13.8	1.5
18–24	55	1.1	0.9	54	0	0
25–34	62	5.5	2	73	1.1	0.8
35–44	52	13.7	3.3	94	5.2	1.5
45–54	114	37.5	2.4	116	15.2	1.8
55–64	74	41.4	3.1	97	35.6	2.6
65–74	41	46.3	4.2	64	49.8	3.3
Puerto Ricans						
18–74	495	15.6	1.1	835	11.5	0.7
18–74 age-adjusted		19.7	1.8		18	1.3
18–24	112	1.1	0.7	188	0.6	0.4
25–34	107	4.6	1.6	171	1.9	0.8
35–44	73	17.4	3.6	152	7.1	1.6
45–54	102	26.4	2.5	174	23.7	1.8
55–64	78	51.1	3.2	97	47	2.7
65–74	23	**	**	53	55.2	3.8

SOURCE: NCHS, HHANES, 1982–84

*Defined as the average of two blood pressure measurements >= 140/90 mm Hg or currently taking antihypertensive medication.

**Sample size too small to produce reliable estimates.

TABLE 2—Distribution of Blood Pressure Level Regardless of Medication Status

	Mexican American		Cuban American		Puerto Rican	
Blood Pressure Level	Males n=1553 % SE	Females n=1965 % SE	Males n=398 % SE	Females n=498 % SE	Males n=495 % SE	Females n=835 % SE
Normal Blood Pressure*	72.7 ± 1.4	85 ± 0.9	67.8 ± 1.4	81.4 ± 1	78.7 ± 1.3	88.2 ± 0.8
High Normal Blood Pressure*	12.3 ± 1	4.9 ± 0.5	13 ± 1	6.6 ± 0.6	6.9 ± 0.7	3.7 ± 0.4
Mild High Blood Pressure	8.8 ± 0.9	4.1 ± 0.5	11.1 ± 0.9	4.8 ± 0.6	10.2 ± 0.9	2.7 ± 0.4
Moderate High Blood Pressure	1.0 ± 0.3	0.2 ± 0.1	1.7 ± 0.4	0.7 ± 0.2	0.3 ± 0.2	0.4 ± 0.1
Severe High Blood Pressure	0.2 ± 0.1	0.1 ± 0.1	0.7 ± 0.2	0 ± 0	0.1 ± 0.1	0 ± 0
Borderline Isolated Systolic Hypertension	4.7 ± 0.7	5 ± 0.5	4.3 ± 0.6	5.3 ± 0.6	2.8 ± 0.5	4.2 ± 0.5
Isolated Systolic Hypertension	0.4 ± 0.2	0.6 ± 0.2	1.2 ± 0.3	1.2 ± 0.3	0.9 ± 0.3	0.6 ± 0.3

*Includes those hypertensives who are controlling their blood pressures.
SOURCE: NCHS, HHANES, 1982–84

Results of comparisons of the distribution of blood pressure levels (regardless of medication status) indicate that Whites or Blacks in NHANES II generally had higher levels in each elevated blood pressure category when compared to Hispanics. Whites and Blacks (both sexes) were two or three times more likely to have mildly elevated diastolic blood pressure (DBP 90-104) than each of the Hispanic subgroups. Among White men and women, as found in the JNC 1984 report, 18.4 percent and 10.8 percent, respectively had mildly elevated diastolic blood pressure. Mildly elevated diastolic blood pressure has been found in 20.6 percent and 17.9 percent of Black men and women, respectively. Whites and Blacks were also more likely to have moderately (DBP 105-114) and severely elevated (DBP > = 115) blood pressures, compared to Hispanics.

The proportion of hypertensives (160/95) aware, treated, and controlled is presented by sex and Hispanic subgroup in Table 5. Awareness was high among women, over 95 percent in each of the three groups. Within each Hispanic subgroup, women hypertensives reported having been told about their hypertension more often than men. A substantial portion of Mexican American and Puerto Rican male hypertensives, i.e. 32.1 percent and 21.3 percent, respectively, are unaware of their condition.

Similarly, a larger number of Mexican American and Puerto Rican men were on no medication for their condition, compared to women in those subgroups. Cuban American hypertensive men were more likely to be on medication than hypertensive men in the other two Hispanic subgroups.

Control rates among hypertensives using the 1990 Objectives definition were generally low. With the hypertension threshold at 160/95 and control threshold at 140/90, the proportions of male hypertensives under control were 22.6 percent, 26.1 percent, and 11.9 percent for Mexican Americans, Cuban Americans, and Puerto Ricans, respectively (see Table 5). Higher rates of control were found among hypertensive women as compared to men, except in the Cuban subgroup where the level of control was similar for male and female hypertensives. No group attained the 1990 Objectives of 60 percent controlled.

TABLE 3—Percent of Individual Blood Pressure Readings (systolic and diastolic) around Critical Values for Definition of Hypertension

Diastolic Blood Pressure	Percent	Systolic Blood Pressure	Percent
86	30.8	136	24.6
88	31.7	138	15.4
90	17.5	140	25.4
92	10.8	142	21.4
94	9.1	144	13.3

SOURCE: NCHS, HHANES 1982–84

Using a less stringent threshold for control (160/95), men still fall short of the objective of 60 percent control; 38.4 percent, 50.6 percent, and 19.7 percent of the hypertensive males (160/95) were controlled for Mexican Americans, Cuban Americans, and Puerto Ricans, respectively.

Discussion

The major finding of this study is that the HHANES populations had significantly lower proportions of hypertensives than did Whites or Blacks in NHANES II. This finding is supported by studies showing lower cardiovascular mortality rates among Mexican Americans and Puerto Ricans.[4,15] However, cardiovascular diseases remain a leading cause of morbidity and mortality in these groups and the importance of hypertension should not be underestimated.[16]

Lower rates of hypertension are contrary to the expectations suggested by levels of risk factors for the condition among Hispanics. Considering the generally lower socioeconomic status of Hispanics,[2] the high prevalence of obesity,[17] and the high prevalence of diabetes,[18] we would anticipate hypertension rates to be higher than Whites. A study of risk factors for hypertension using HHANES data may shed light on this problem.

TABLE 4—Hypertension Prevalence Estimates NHANES II and HHANES (ages 18–74 years)

	Male		Females	
	%	SE	%	SE
*NHANES II**				
White	32.6	1.4	25.3	0.9
Black	37.9	2.2	38.6	1.8
*HHANES***				
Mexican American	22.9	1.1	19.7	0.9
Cuban American	20.5	2	13.8	1.5
Puerto Rican	19.7	1.8	18.0	1.3

SOURCE: NHANES II, 1976–80 (JNC 1984) NCHS, HHANES, 1982–84
*JNC 1984
**Age-adjusted to 1978 population

TABLE 5—Hypertension* Awareness, Treatment, and Control Status among Three Hispanics Subgroups**

	Males		Female		Female/ male % difference	(95% CI)
	%	SE	%	SE		
Mexican American						
Awareness	67.9	5.0	95.0	1.8	27.1	(15.8, 38.4)
Treated	49.3	5.3	86.0	2.9	36.7	(23.9, 49.5)
Controlled	22.6	4.4	43.6	4.1	21.0	(8.3, 33.7)
Cuban Americans						
Awareness	87.2	2.6	95.1	1.6	7.9	(1.4, 14.4)
Treated	70.1	3.6	79.1	3.0	9.0	(−0.9, 18.9)
Controlled	26.1	3.5	29.5	3.3	3.4	(−6.8, 13.6)
Puerto Ricans						
Awareness	78.7	3.8	96.1	1.2	17.4	(9, 25.8)
Treated	41.0	4.6	85.5	2.2	44.5	(33.7, 55.3)
Controlled	11.9	3.0	41.6	3.1	29.7	(20.6, 38.8)

SOURCE: NCHS, HHANES, 1982–84

*Hypertension is defined as either >=160/95 or reported currently taking antihypertensive medication.

**Control is defined as a threshold of 140/90.

Previous surveys comparing blood pressure levels of Hispanics with Whites have given inconclusive results. The results of the San Antonio Heart Study, conducted between 1979–82, suggests that Mexican American males had a prevalence of hypertension approximately equal to Whites, and Mexican American females had lower rates than White females.[4] In Orange County, California, Hispanics and non-Hispanics were found to have similar systolic and diastolic blood pressure.[5] The California Hypertension Survey (1978) found that Hispanic men had slightly lower rates of hypertension than White men. Among women, the prevalence of hypertension was higher in Hispanic women than in White women in the younger age groups and somewhat lower in older age groups.[6] The one study of blood pressure among Puerto Ricans was small in scale and nonrepresentative.[7] No study of Cuban American blood pressure has previously been published. The conclusions of these earlier studies which differ from those presented here may be due to the larger and more representative sample selected in HHANES.

The limitations of these data include the facts that: 1) hypertensives controlling their blood pressure without medication are not included in the definition; 2) only the civilian, non-institutionalized population ages 18–74 are included; 3) the sampling universe defined in HHANES does not include all persons of Hispanic origin; and 4) measures were taken on only one occasion. The use of blood pressure measurements on one occasion limits both the definition of hypertension and the "1990's" definition of control. Clinical definitions require more than one elevated measurement and the "1990 Objectives" define long-term control as "at or below 140/90 for two or more years."

The comparison of the prevalence of hypertension among Whites, Blacks, and Hispanics presented here must address the possibility of temporal changes. The NHANES II and HHANES were conducted five years apart. The prevalence of hypertension in the general population has remained relatively stable over the past 20 years.[19] The incidence of hypertension also seems to have been stable for this period as measured in the Framingham study population.[20] Comparisons between the temporally separated NHANES II and HHANES are useful if we reasonably assume the rates among Hispanics have also been stable.

No comparisons of Whites or Blacks with the three Hispanic subgroups were attempted for awareness, treatment, and control status. Direct comparison of HHANES and NHANES II for the status of awareness, treatment, and control of hypertension is inappropriate because of well described secular trends. Over the past decade, awareness of hypertension has increased, more people have begun treatment, and control has improved.[21–23] These important changes were happening during the interval between NHANES II (1976–80) and HHANES (1982–84). To the extent that US Hispanics have participated in the transition, we would expect that awareness, treatment, and control of hypertension also would have improved since the time when NHANES II was conducted.

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APPENDIX

Diastolic Blood Pressure (DBP) (mm Hg)	Category*
<85	Normal Blood Pressure
85–89	High Normal Blood Pressure
90–104	Mild Hypertension
105–114	Moderate Hypertension
≥115	Severe Hypertension
Systolic Blood Pressure (SBP) (mm Hg) when DBP <90 mm Hg	Category*
<140	Normal Blood Pressure
140–159	Borderline Isolated Systolic Hypertension
≥160	Isolated Systolic Hypertension

*A classification of borderline isolated systolic hypertension (SBP 140 to 159 mm Hg) or isolated systolic hypertension (SBP ≥ 160 mm Hg) takes precedence over a classification of high normal blood pressure (DBP 85 to 89 mm Hg) when both occur in the same individuals. A classification of high normal blood pressure (DBP 85 to 89 mm Hg) takes precedence over a classification of normal blood pressure (SBP <140 mm Hg) when both occur in the same individual.

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IOM Conference on Health Status Assessment: Call for Papers

The Institute of Medicine of the National Academy of Sciences has announced a call for papers for its September 1991 conference on Advances in Health Status Assessment. Building on the outcome of two previous conferences (1986 and 1988) this conference will concentrate on the use and usefulness of health status and health-related quality of life measures in clinical practice, especially ambulatory care, and in clinical research that will directly facilitate the application of these measures by practitioners. The following issues are of particular interest, although others will be considered, and empirical research papers are especially invited.

- The use of measures of health status and health-related quality of life as aids in screening, planning patient care, monitoring individual patient outcomes of care, or evaluating quality of care in ambulatory clinical settings;
- Improvements in the definition, reliability, validity, and other methodologic or statistical aspects of health status and outcome measures, as they related to clinical settings.

The deadline for submission of abstracts of at least 2500 words is February 15. Author notification will be on or about April 1, with first drafts due June 1 and final publishable versions due August 1. For further information, contact Kathleen N. Lohr, PhD, or Jo Harris-Wehling, MPA, IOM, NAS, 2101 Constitution Avenue, NW, Washington, DC 20418; (202) 334-2165.

A Note on the Measurement of Hypertension in HHANES

Arline T. Geronimus, ScD, Lisa J. Neidert, PhD, and John Bound, PhD

Abstract: Using data from the HHANES, we found the rates of elevated blood pressure readings on clinical examination to be extremely low for a sample of Mexican American and Puerto Rican women. The prevalence rates were one-fourth to one-fifth the rates found for a comparable sample of White women from NHANESII. These findings are discrepant with the little that is known about hypertension prevalence among Hispanics and with estimates of hypertension prevalence for Mexican Americans and Puerto Ricans drawn from NHANESII. While our HHANES samples women had much lower rates of clinical high blood pressure than Whites, they reported hypertension histories in excess of Whites. Rates of medicine usage among Hispanics were insufficiently large for effective treatment to explain the disparity. The prevalence estimates increased, but the relative discrepancies remained when we altered our sample specifications and clinical high blood pressure measure. A possible explanation for these discrepancies is that few physicians performed the majority of blood pressure readings in our HHANES sample. This may have been statistically inefficient. The discrepancies noted suggest that HHANES may not be a reliable source of information on hypertension among Hispanic women. (*Am J Public Health* 1990; 80:1437–1442.)

Introduction

Hypertension is one of the leading causes of death in the United States. It is an illness that can be easily diagnosed and in the aggregate has responded well to medication. Over the past 20 years, there has been an increase in the proportion of the hypertensive population diagnosed as having hypertension, an increase in the proportion using antihypertensive medicine, and a decrease in the incidence of clinical hypertension.[1] Despite the improvements in the diagnosis, treatment, and control of hypertension in the United States, it remains a persistent health problem, especially for Blacks.[1–3]

Less can be said about the historical or current trends in hypertension for the various Hispanic populations. Most available data have been local area, small-sample studies[4–8] which cannot be generalized to the national population. While Hispanic respondents have often been identified in national health surveys, their numbers were usually too small to analyze well. Often, the definition of the Hispanic population is an aggregation of all Spanish-speaking groups, a practice which ignores the real differences among the various Hispanic groups.[9–11]

With the release of data from the 1982-84 Hispanic Health and Nutrition Examination Survey (HHANES), the potential to apply sophisticated analytical approaches to understanding disease patterns within Hispanic populations is enhanced. HHANES is the only survey that has both large representative samples of members of Hispanic groups within the United States and includes objective measures of health through the examination data, rather than relying only on self-reports. Thus, the validity of the clinical data is critical for the research and policy questions HHANES is intended to address. However, as this paper suggests, there may be problems with the HHANES clinical blood pressure data for women.

Methods

The focus of this analysis is the hypertension data in HHANES. However, to provide a comparison population, information from the National Health and Nutrition Examination Survey, 1976–80 (NHANESII) is also used. Both the NHANESII and HHANES data sets utilized complex, multistage, stratified, clustered samples of defined populations. HHANES differs from NHANESII and previous health interview surveys conducted by the National Center for Health Statistics in that it was not a national probability survey, but a survey of three special subgroups of the population who reside in three geographically distinct regions: Mexican Americans in the southwest United States, Puerto Ricans in the New York area, and Cuban Americans in Miami, Florida. Like NHANESII, the data include an interview and a direct medical examination for a sample of civilian, noninstitutionalized persons 6 months to 74 years of age. (For more details on the sampling procedures used in HHANES and NHANESII, see the NCHS plans for operation.[12,13])

Our ongoing work on the health status of women in their childbearing years led us to the discovery of potential problems with the HHANES blood pressure data. Our sample included all currently menstruating women, who were not pregnant at the time of the survey, and were between the ages of 12 and 54 years. The White and Black members of our sample were drawn from NHANESII while the Puerto Rican and Mexican American sample members were drawn from HHANES. Only respondents who identified themselves as Mexican in the southwest sample or Puerto Rican in the northeast sample were used in the analysis. The resulting sample sizes were 1,700 Mexican Americans, 757 Puerto Ricans, 2,963 Whites, and 509 Blacks. Persons of Hispanic origin have been excluded from the White and Black populations. Due to small numbers, Cuban American women were excluded.

For the HHANES examinations, blood pressure measurements were obtained using procedures recommended by the American Heart Association. For the Black and White populations (NHANESII), three readings were taken (sitting, recumbent, sitting), while for the Mexican American and Puerto Rican populations (HHANES), two readings were taken (both sitting). Additional questions on hypertension history, treatment history and compliance, and current medicine usage were ascertained through the detailed medical history interview. Averages of the available blood pressure

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Address reprint requests to Arline T. Geronimus, ScD, Assistant Professor, Department of Public Health Policy and Administration, School of Public Health, University of Michigan, 1420 Washington Heights, Ann Arbor, MI 48109–2029. Dr. Neidert is Research Associate at the School of Public Health and the Population Studies Center; Dr. Bound is Assistant Professor, Department of Economics, and Research Associate, Population Studies Center, both also at the University of Michigan. This paper, submitted to the Journal April 27, 1989, was revised and accepted for publication April 5, 1990.

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readings and some of the interview information provide the basis for our determination of hypertension status. The following indicators of hypertension are used in the analysis:

- Respondent has clinical high blood pressure (HBP). In our initial sample clinical HBP was defined as mean arterial pressure (MAP) greater than 105 where MAP = [systolic + (2 × diastolic)]/3.[14,15] As explained below in the text, in later specifications we defined it as systolic ≥ 140 of diastolic ≥ 90.
- Respondent is currently taking antihypertensive medicine.
- Respondent has ever been told by a doctor that she has high blood pressure or hypertension.

It is standard procedure to determine hypertension status on the basis of the first two of these three indicators.[1] Thus, we initially combined them to identify hypertensives. However, when the clinical data appeared questionable, we also looked at the third indicator to learn if the history and clinical data were consistent.

Standard errors are calculated under the assumption that the NHANESII data are simple random samples of the population. The complex sample design implies that these standard errors are almost certainly too small, by a factor that could be as large as $\sqrt{2}$. Initially we attempted to account for the complex samples design using both Taylor series expansion and balanced half sample methods.[16-20] Estimated design effects for the Black, Mexican American, and Puerto Rican samples were often implausibly low (less than 1). These results are consistent with those found by NCHS staff.[21,22] Following NCHS suggested guidelines,[23] which involve averaging design effects across subsamples, did not yield more reasonable results. We thus thought it most reasonable to simply report standard errors without taking into account the complex sample design while warning the reader to interpret results cautiously.

Using the direct method,[24] summary statistics for Blacks, Mexican Americans, and Puerto Ricans are standardized to the age distribution of our White sample.

Results

HHANES Clinical High Blood Pressure Data

Tables 1 and 2 provide selected measures of hypertension according to race and Hispanic ethnicity. Not only do Mexican Americans and Puerto Ricans have the lowest prevalence of elevated high blood pressure readings (clinical HBP), but the overall rate for Mexican American women is only one-third the White rate (1.7 versus 5.1) while the Puerto Rican rate is not even one-fourth of the White rate (1.2 versus 5.1). Even allowing for large design effects, these differences are clearly significant. The rates of clinical HBP also look questionable when they are examined by age (Table 2). Among Puerto Rican women, there are no cases of high blood pressure under the age of 35. Among Mexican American women there are hypertensives in all but the youngest group (<15) but, unlike Black or White women, even at the oldest ages hypertension prevalence does not approach double digits.

TABLE 2—Prevalence of Clinical High Blood Pressure by Race/Ethnicity and Age

	Hispanic		Non-Hispanic	
Age (years)	Mexican American	Puerto Rican	White	Black
<15	0.0	0.0	0.7	0.0
15–24	0.1	0.0	1.1	1.7
25–34	2.0	0.0	3.1	7.3
35–44	3.3	3.1	9.7	22.8
45–54	5.0	9.8	21.9	48.4
Total[a]	1.7	1.2	5.1	9.2

[a]Standardized to the White non-Hispanic population
SOURCE: HHANES, 1982–1984 and NHANESII, 1976–1980.

The low magnitudes and unusual age patterns of hypertension prevalence we observed in HHANES based on clinical high blood pressure were unanticipated. That both Hispanic populations had prevalence rates that were not simply lower than, but were actually only a fraction of, the White was unexpected. This led us to question whether Mexican American and Puerto Rican women do indeed have extremely low rates of hypertension or whether the clinical high blood pressure data for Puerto Rican and Mexican American women in HHANES were downwardly biased.

To examine the matter further we conducted the following exercises: 1) we compared our findings to those in the literature on high blood pressure among Mexican Americans and Puerto Ricans; 2) we compared the clinical high blood pressure data in HHANES to that reported for the Mexican American and Puerto Rican respondents in NHANESII; 3) we considered other measures of hypertension in HHANES; 4) we checked the sensitivity of our prevalence estimates to our sample specifications and clinical measure of high blood pressure; and 5) we considered the possibility that study effects, specifically physician effects, influenced the HHANES clinical data.

Comparisons with Other Studies

A few recent studies have examined hypertension in the Mexican American and Puerto Rican populations.[4-8] The one Puerto Rican study was of a highly selected sample in the Bronx (participants at a health fair), from which it is difficult to generalize. However, hypertension rates were very high, with the majority of sample members hypertensive. Most studies of Mexican Americans have been local area studies and, thus, cannot be generalized to the US Mexican population. These local area studies have not usually focused

TABLE 1—Selected Measures of Hypertension by Hispanic Ethnicity and Race, Age Standardized to the White Non-Hispanic Population, with Standard Errors

	Clinical HBP				Current Medicine Usage				HBP History				Clinical HBP + Medicine Usage			
	Mx	PR	Wh	Bl	Mx	PR	Wh	Bl	Mx	PR	Wh	Bl	Mx	PR	Wh	Bl
Percent Hypertensive (Standard Error)	1.7 (.3)	1.2 (.4)	5.1 (.4)	9.2 (1.3)	2.2 (.4)	2.1 (.5)	2.7 (.3)	5.2 (1.0)	13.7 (.8)	17.2 (1.4)	12.8 (.6)	19.3 (1.7)	3.5 (.4)	2.8 (.6)	6.5 (.4)	12.2 (1.4)

SOURCE: Menstruating, non-pregnant women ages 12-54 HHANES, 1982-84 and NHANESII, 1976-80 specifically on women and, when women have been included, the age stratification is not as detailed as in our study. Nevertheless, data from the local area studies indicate that the Mexican Americans do not have extremely low rates of hypertension. For example, both a Laredo, Texas[4] and a Sacramento, California[7] study found rates of hypertension for Mexican Americans to be intermediate between Whites and Blacks, while a study comparing Mexican Americans and non-Hispanic Whites in San Antonio, Texas found them to have roughly comparable rates.[6] Point estimates were slightly lower for Mexican American women compared to non-Hispanic Whites, but this difference was not statistically significant. We were able to simulate a portion of the Laredo, Texas study using the HHANES data. Using the same sample specifications and the same measure of clinical hypertension as used in that study (diastolic $\geq$ 95), we were able to calculate a prevalence estimate for one age group, women 40–49 years. While in the Laredo study the point estimate for this group was 7.9 percent, using HHANES data, ours was only 1.9 percent.

Hispanics in HHANES versus NHANESII

Using the NHANESII data, Table 3 presents several measures of hypertension for Mexican American, Puerto Rican, Black, and White women. It is important to note sample sizes. There are only 35 Puerto Rican women in NHANESII, too few for analysis of a rare event such as hypertension. There are 140 Mexican American women which is much less than what one could find in HHANES, but enough for rough comparisons.

Unlike in HHANES, Hispanic hypertension rates in NHANESII are fairly comparable to those found among White women. The point estimates for Hispanic women in NHANESII are actually higher than those for Whites. Using NHANESII, the ordering of prevalence rates by racial/Hispanic ethnic identification is what one might expect based on the local area studies and what is more generally known about the relative health and socioeconomic status of US Black, Whites, Puerto Ricans and Mexican Americans.[25]

Other Measures of Hypertension in HHANES

HHANES has several indicators of hypertension besides the clinical measure, most prominently self-reported history data. The clinical data and the hypertension history data in HHANES appear discrepant (see Table 1). If Mexican American women had a low prevalence of hypertension, we might expect to see a low prevalence of positive hypertension histories. Mexican American and White women report positive hypertension histories in roughly equal proportions, while Puerto Rican women report a positive hypertension history significantly *more* often. Inspecting the data within age groups, we observed additional anomalies. Examples are (not reported in the tables): among the 27 Puerto Rican women 48 to 50 years of age, 13 (43 percent) report a history of high blood pressure; yet only two currently have an elevated clinical reading and three are taking antihypertensive medicine. Among the 42 Mexican American women in this age interval, 15 (33 percent) report a history of high blood pressure; yet only nine are defined as currently hypertensive (four with clinical high blood pressure and five currently using antihypertensive medication). Some of the "history only" women were probably reporting an experience with a finite hypertensive period such as pregnancy-induced hypertension. However, to explain the entire discrepancy by this fact, one would have to assume that Mexican American and Puerto Rican women are far more prone to such periods than are Black or White women, since discrepancies of this magnitude were not found among Blacks or Whites. Similarly, to infer recall bias explains the entire difference would imply that Hispanics are more likely to misreport than Blacks or Whites, and that both Hispanic groups misreport in the same direction. This seems unlikely and is also unsupported by the Hispanic data in NHANESII. Among Mexican Americans and Puerto Ricans in NHANESII, there is a −0.5 point and a 3.5 point difference, respectively, between the percent reporting a history of hypertension and those with evidence of current hypertension. The differences in the HHANES data are much greater at 10.2 points (Mexican Americans) and 14.4 points (Puerto Ricans).

We also considered the effects relevant secular trends might have on our findings. The NHANESII and HHANES surveys were implemented several years apart. During this time efforts were made to educate the public about the effects of hypertension and the importance of treatment. The secular trends in awareness would lead us to expect to find higher percentages reporting a history of hypertension and higher percentages reporting experience with antihypertensive treatment in HHANES compared to NHANESII, but hypertension prevalence estimated as the combination of elevated readings and medicine usage would not be expected to be lowered dramatically. Comparing the HHANES and NHANESII results, the proportions of self-reported histories did increase between the two surveys, as expected, but hypertension prevalence was lowered far beyond what one might expect. The rate of elevated clinical readings decreased 75 percent for Mexican Americans and 86 percent for Puerto Ricans; the prevalence rate (elevated readings plus medicine usage) decreased 54 percent for Mexican Americans and 68 percent for Puerto Ricans.

Questions are also raised by comparisons of the medicine usage and hypertension prevalence data. If the clinical data are taken at face value, the figures in Table 1 show that 63 percent of hypertensive Mexican American women and 75 percent of hypertensive Puerto Rican women take medicine. The corresponding figure for White women is 42 percent and for Black is 43 percent. These figures suggest that hypertensive Mexican American and Puerto Rican women (women who are members of populations usually described as medically underserved) undergo medical treatment in higher

TABLE 3—Selected Measures of Hypertension for Mexican American and Puerto Rican Women: Data from NHANESII

	Sample Size				Clinical HBP		HBP History		Clinical HBP + Medicine Usage			
	Mx	PR	Wh	Bl	Mx	PR	Mx	PR	Mx	PR	Wh	Bl
Total[a] (Standard errors)	140	35	2693	509	6.9 (2.1)	8.7 (4.8)	7.1 (2.2)	12.2 (5.5)	7.6 (2.2)	8.7 (4.8)	6.5 (.4)	12.2 (1.4)

[a] Standardized to the White non-Hispanic population.
SOURCE: NHANESII, 1976–80.

proportions than hypertensive members of other populations. An alternative explanation for these percentages would be that they are artificially inflated by a downwardly biased denominator, i.e. that the clinical data are inaccurate.

Altering Sample Specifications

We wondered to what degree the estimates we calculated were sensitive to our specific sample specifications and measure of high blood pressure. Consistent with the aims of our larger research program, we had restricted our sample to menstruating women and had used a measure of high blood pressure (MAP) that is commonly used in the obstetric/gynecological literature[26,27], but is not as common in other areas of clinical practice or research.

In Table 4, we report hypertension prevalence estimates (systolic ≥ 140 or diastolic ≥ 90) by age for all Mexican American and Puerto Rican women in HHANES, ages 12–54. Data on medicine usage, history, and a measure that combines clinical and medicine usage data are also reported for these women, as are data for Whites, Blacks and Hispanics in NHANESII. With these changes, we expected rate increases, because MAP is a conservative measure of hypertension that excludes some borderline hypertensives. As expected, the new estimates are higher than those using our previous sample specifications and measure, especially at the older ages where menopausal women have been added. However, the contrasts between clinical hypertension status versus hypertension history remain large. The clinical rates for Puerto Rican and Mexican American women in HHANES remain much lower than the White or Black rates, in each case being no more than one-half the White rate. Again, applying the new specifications to data from NHANESII suggests a different picture from HHANES, suggesting that our initial findings were not simply an artifact of the original specifications.

A Possible Explanation

Great effort was put into quality control when the HHANES data were collected, but our findings raise the question of whether some sort of study effect might explain the low rates of elevated blood pressure readings among examinees. One possibility is the fact that only a few physicians were relied upon to perform the majority of the clinical examinations. This practice has the potential to influence the validity of hypertension data. In general, blood pressure readings have been observed to be subject to variation between examiners.[28,29] This is not to suggest that some physicians' readings are incorrect, but only that there is some expected variation in reporting around "true scores." If there were many physicians evenly distributed across cases, the physician effects would tend to cancel out. However, a reliance on a small number of physicians introduces an extra element of sampling error, one that will not necessarily be accounted for by the calculation of standard errors.[30] As the National Center for Health Statistics has, itself, warned in the past, this practice can seriously jeopardize the validity of a survey.[29]

Table 5 shows the distribution of physicians across the Mexican American and Puerto Rican surveys in our initial sample. Each survey used only four physicians. The distribution of these examiners is uneven, particularly in the Mexican American sample. In the Mexican American survey, Doctors 501 and 502 examined 96 percent of the cases. In the Puerto Rican sample, the examinations are spread across three doctors who, as a group, performed 98.7 percent of the readings.

As shown in the Appendix, we performed sample calculations to estimate the potential statistical inefficiency that might have resulted from the usage of so few examining physicians in HHANES (effectively two for the Mexican American sample and three for the Puerto Rican). The magnitude of the inefficiency produced by using so few physicians depends on both the ratio of the sample size to the number of physicians and on the magnitude of the physician effects. Our estimates suggest that standard errors could have been enlarged three-fold for Puerto Ricans and six-fold for Mexican Americans by the employment of so few physicians.

We conclude from these sample calculations that there appears to be enough variability between examining physicians in the proportion they are likely to diagnose as hypertensive that using only two or three physicians has the potential to reduce substantially the reliability of the survey. However, we present this only as one possible explanation for the low prevalence estimates we found. The data do not permit us to test or rule out other possible systematic biases. We cannot know how likely the physician effects explanation is relative to alternative explanations based on other potential sources of systematic bias.

TABLE 4—Selected Measures of Hypertension by Race and Ethnicity (Clinical HPB Defined as Systolic ≥140 or Diastolic ≥90) All Women in HHANES or NHANESII

	Mexican American				Puerto Rican				White				Black			
Age (years)	CI[a]	CI+M	M[b]	Hx[c]	CI	CI+M	M	Hx	CI	CI+M	M	Hx	CI	CI+M	M	Hx
<15	0.0	0.0	0.0	1.7	0.0	2.5	2.5	3.3	2.9	2.9	0.0	2.0	3.2	3.2	0.0	2.4
15–24	0.3	0.9	0.6	8.1	0.0	0.0	0.0	8.8	2.7	3.1	0.3	6.0	6.3	7.3	1.1	11.5
25–34	3.5	5.2	2.6	17.7	0.5	2.0	1.5	18.3	4.3	5.5	1.5	13.3	12.6	15.1	3.7	18.8
35–44	4.9	7.5	4.2	18.1	5.0	7.2	2.7	26.5	15.2	17.4	5.0	21.0	30.9	37.5	13.9	39.7
45–54	17.6	24.1	12.6	35.1	17.2	24.1	15.4	42.6	28.9	36.1	15.4	29.6	52.6	66.1	40.2	60.2
Totals:																
HHANES	5.2	7.5	4.0	16.9	4.3	6.6	4.0	20.4								
(Standard Errors)	(.5)	(.6)	(.4)	(.8)	(.7)	(.1)	(.7)	(1.3)								
NHANESII	12.1	13.2	2.8	11.2	13.2	13.2	4.3	12.7	10.4	12.6	4.3	14.7	21.0	25.7	11.4	27.1
(Standard Errors)	(2.6)	(2.7)	(1.3)	(2.5)	(4.9)	(4.9)	(3.0)	(4.9)	(.5)	(.5)	(.3)	(.6)	(1.6)	(1.7)	(1.3)	(1.8)

a)CI = Clinical High Blood Pressure
b)M = Antihypertensive Medicine Usage
c)Hx = History of Hypertension

TABLE 5—Distribution of Doctors According to Sample

Doctor	Sample	
	Mexican American	Puerto Rican
500	3.4	0.0
501	54.9	0.0
502	41.5	40.2
504	0.0	21.2
505	0.0	37.3
510	0.2	1.3
Sample Size	1700	757

SOURCE: HHANES, 1982–84

Conclusions

The several features of the HHANES hypertension data for women we have described suggest that it is wise to be wary of these data and exercise caution when interpreting them. The prevalence of elevated readings at the HHANES examination is low compared to analyses using other data. These examination data appear more discrepant with the history data than is true in NHANESII or that can be accounted for reasonably by recall bias. The validity of the HHANES clinical high blood pressure data may be threatened by the reliance on only a few examining physicians.

Hispanic groups in the US are understudied populations. Given this lack of benchmark data to which we would compare the HHANES data under ideal circumstances, we cannot conclude definitively that the low rates of elevated blood pressure readings estimated for reproductive-age women in HHANES are invalid. However, careful review of the available information reveals inconsistencies that suggest to us that hypertension prevalence estimates based on the HHANES examination data are downwardly biased.

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APPENDIX

The magnitude of the inefficiency due to the reliance on few examining physicians depends on both the ratio of the sample size to the number of doctors and on the magnitude of the doctor effects. In particular, imagine that n doctors each examine m respondents. Also, let the true population proportion with hypertension be π, while the probability that the i_{th} doctor identifies a person as hypertensive is π_i. While we might assume that

$$E_i(\pi_i) = \pi,$$

the fact that there are doctor effects implies that, in general, $\pi_i \neq \pi$. A natural measure of the magnitude of these effects is just the variance of the π_i's,

$$\sigma^2_{\pi_i} = E_i(\pi_i - \pi)^2.$$

Were it not for the doctor effects the sampling variance of p, the proportion of respondents with hypertension, on a sample of size $N = mn$ would just be

$$\frac{\pi(1 - \pi)}{N}.$$

Doctor effects add an extra component to this variance. With doctor effects the variance will be

$$\frac{E_i[\pi_i(1 - \pi_i)]}{N} + \frac{\sigma^2_{\pi_i}}{n}.$$

Thus the efficiency of the survey design relative to that of a simple random sample is just:

$$\frac{\dfrac{E_i[\pi_i(1 - \pi_i)]}{N} + \dfrac{\sigma^2_{\pi_i}}{n}}{\dfrac{\pi(1 - \pi)}{N}}$$

Note if either the number of doctors used, n, is larger or doctor variability, $\sigma^2_{\pi_i}$ is small the above expression approximates 1 implying that in these circumstances the inefficiency is small.

To evaluate this expression we need an estimate of $E_i[\pi_i(1 - \pi_i)]$ and $\sigma^2_{\pi_i}$. With effectively only two doctors in the Mexican and three in the Puerto Rican samples there are too few doctors to allow a reasonable estimate of $\sigma^2_{\pi_i}$. What we can do, though, is use the sample of Whites in NHANESII to calculate sample analogies for the above expressions. From these we can estimate the impact that using only two or three doctors would have on the efficiency of the NHANESII survey.

In the NHANESII survey the 2,963 White women were surveyed by a total of 13 doctors. The sample variance,

$$S^2_{\pi_i} = \frac{\sum_i(p_i - \bar{p}_i)^2}{12}$$

is .00230. The sample variance reflects both systemic differences between the doctors and sampling variability. To get an estimate of $\sigma^2_{\pi_i}$ we need to subtract off from $S^2_{\pi_i}$ an estimate of the sampling variability. This we can estimate as

$$\frac{\sum_i\left[\dfrac{p_i(1 - p_i)}{m_i}\right]}{n}$$

= .00046. Thus a calculation gives us an estimated $\sigma^2_{\pi_i}$ of .00230 − .00046 = .00184. To estimate $E_i[\pi(1 - \pi)]$ we simply use

$$\frac{\sum_i p_i(1 - p_i)}{12}$$

= .04580.

Now suppose that we were to redo the NHANESII survey using only two doctors, each of whom tests 800 respondents (similar to the distribution in the Mexican sample). Evaluating the above formula implies that the efficiency of the random sample is 35 times that of the sample using only two doctors. In other words, the standard errors are almost six times what they would have been with random sampling. Redoing the calculation assuming that three doctors each do 250 cases (similar to the distribution in the Puerto Rican sample), we find that a random sample with the number of respondents would be roughly 11 times as efficient or the standard errors are more than three times as large as they would be with a random sample. We conclude from these samples calculations that there appears to be enough variability between physicians in the proportion they are likely to diagnose as hypertensive that relying on only two physicians to examine the Mexican American women and only three for the Puerto Rican women has the potential to reduce substantially the survey's reliability.

Fruit and Vegetables in the American Diet: Data from the NHANES II Survey

Blossom H. Patterson, MA, Gladys Block, PhD, William F. Rosenberger, MPhil,
David Pee, MPhil, and Lisa L. Kahle, BA

Abstract: Twenty-four hour dietary recall data from the Second National Health and Nutrition Examination Survey (1976–80) were used to estimate the numbers of servings of fruit and vegetables consumed by Black and White adults, to examine the types of servings (e.g., potatoes, garden vegetables, fruit, and juice), and to estimate the mean intake of calories, fat, dietary fiber, and vitamins A and C by number of servings. An estimated 45 percent of the population had no servings of fruit or juice and 22 percent had no servings of a vegetable on the recall day. Only 27 percent consumed the three or more servings of vegetables and 29 percent had the two or more servings of fruit recommended by the US Departments of Agriculture and of Health and Human Services; 9 percent had both. Consumption was lower among Blacks than Whites. The choice of vegetables lacked variety. Diets including at least three servings of vegetables and two servings of fruit contained about 17 grams of dietary fiber. Although caloric and fat intake increased with increasing servings of fruit and vegetables, the percent of calories from fat remained relatively constant. Although these data are 10 years old, more recent surveys have shown similar results. The discrepancy between dietary guidelines and the actual diet suggests a need for extensive public education. *(Am J Public Health* 1990; 80:1443–1449)

Introduction

In a recent article,[1] it was found that a large proportion of the adult United States population eats no vegetables (0.17) or fruit (0.41) on any given day. The Surgeon General's Report on Nutrition and Health[2] recommends increasing consumption of vegetables and fruit. Based on a comprehensive literature review, the National Academy of Sciences (NAS) concluded that diet influences the risk of several major chronic diseases and recommended eating five or more daily servings of a combination of vegetables and fruit, especially green and yellow vegetables and citrus fruit.[3] The US Department of Agriculture (USDA) and the Department of Health and Human Services (DHHS) recommend as part of their food guidance system that the daily diet include two to three servings of fruit and three to five servings of vegetables.[4,5] These specific recommendations are in the context of general recommendations for a diet that meets nutrient requirements, does not include excessive amounts of fat (in particular saturated fats), and is consistent with maintaining desirable weight. It is important therefore to determine what relation increased fruit and vegetable consumption might have, not only to intake of obviously related nutrients such as vitamins A and C but also to fat intake and maintenance of appropriate body weight. In the NHANES II data, collected between 1976 and 1980, it is possible to examine these relationships in self-selected diets that were actually consumed during a 24-hour period by a representative sample of US adults.

The article mentioned above[1] described the proportions of the US population that consumed any amount of a given food, however large or small; no attempt was made to determine the number of servings consumed. In the present inquiry we estimate the number of servings consumed, taking portion size into account. We also examine the intake of calories, fat, fiber, and vitamins A and C in the diets of persons consuming various numbers of servings of fruit and vegetables. While it is obvious that intake of vitamins A and C would increase with increasing numbers of servings, the relative contribution of fruit versus vegetables is not well known, nor is the nutrient intake actually achieved in self-selected diets. Similarly, the dietary fiber intake contained in diets including multiple servings of fruit and vegetables may be inadequately appreciated by the general public.

Methods

The NHANES II survey was conducted by the National Center for Health Statistics between 1976 and 1980.[6] A highly stratified multi-stage probability design was used to obtain a representative sample of the civilian noninstitutionalized population, ages six months to 74 years; we report on 10,313 White and 1,335 Black adults ages 19 to 74. We excluded other races because of small numbers, as well as imputed, unreliable, or surrogate data. Results are based on weighted data, permitting inference about the total Black and White noninstitutionalized US population.[7] Group means and standard errors were calculated using software appropriate for complex sample survey data.[8] Standard errors are frequently large for Blacks, due to their small number in the survey. Therefore, results for Blacks should be viewed with caution. Conversely, standard errors for Whites are small due to the very large sample size, so that small differences in intake, while not biologically meaningful, could be statistically significant.

Dietary interviews were conducted by interviewers with a knowledge of food preparation and nutrient composition. A measure of portion size was obtained using three-dimensional food models, as well as dishes and glassware of various sizes. Subjects were asked to report all food and drink consumed in the 24-hour period before the interview day. Almost all recall days were weekdays. Food codes, food weight, and nutrient values were based on data from the USDA, industry, and other sources.[9] The coding of these foods has been described elsewhere.[10] Based on reported portion sizes, the number of grams consumed was calculated and included on the 24-hour recall tape[11] for each food reported by each respondent.

Address reprint requests to Blossom H. Patterson, MA, Clinical and Diagnostic Trials Section, Biometry Branch, Division of Cancer Prevention and Control, National Cancer Institute, Executive Plaza North, Rm 344, Bethesda, MD 20892. Dr. Block is also with that Section at NCI; the other co-authors are with Information Management Services, Inc., Silver Spring, MD. This paper, submitted to the Journal August 3, 1989, was revised and accepted for publication May 10, 1990.

We converted the number of grams reported for each fruit and vegetable into a number of servings as follows. We used as a guide suggested serving sizes given by a food guidance system developed to help implement the USDA/DHHS dietary guidelines.[5] For fruits, a serving is defined as an average piece of whole fruit or six ounces of fruit juice, and for vegetables, one-half cup, cooked or raw. An examination of a sample of actual food records showed that some reported amounts of fruit or vegetables, such as a slice of onion on a hamburger, or lemon juice added to tea, were too small to be considered a serving, while others were so large that they should be considered as more than a single serving. We chose a lower limit of one ounce (30 grams) for any serving of a fruit or vegetable, and two ounces (62 grams) for fruit juice. Small lower limits on a single serving were chosen so as not to eliminate portions consumed by persons with small appetites. This choice may result in some overestimation of the proportions eating one or more servings on the recall day. We chose an upper limit for a single serving so that a large portion (eaten on a single occasion) would be treated as more than a single serving. A single portion of a fruit or vegetable was considered to be at most two servings so as not to overestimate the number of servings of big eaters. Juice consumed on a single eating occasion was considered (at most) three servings.

In order to convert a large portion of whole fruit or vegetables into two servings, we determined a representative or "typical" number of grams in a half cup serving of vegetables, and the weight of an "average" piece of fruit, using weights given for the most popular[1] vegetables and fruits in Agriculture Handbook No. 456.[12] A half cup serving of many vegetables weights about 75 grams, and many whole fruits weigh roughly 120 grams. We defined as two servings any portion of vegetables (except salad) weighing 150 grams or more, and any portion of fruit (excluding fruit juice) weighing 240 grams or more. For fruit juice, individuals consuming between 12 and 18 ounces (372 to 557 grams) were given credit for two servings of fruit, and those consuming 18 or more ounces (558 or more grams) were given credit for three servings. An individual eating a green salad on a single occasion was given credit for a single serving rather than for multiple servings corresponding to each ingredient. Thus individuals could be credited with a maximum of one salad per eating occasion, or a maximum of the four different occasions per day coded in the survey.

Individual foods included in each food category are given elsewhere.[1] Vegetables include potatoes, dried peas and beans (e.g., black-eyed peas, kidney beans), green salad (primarily lettuce and raw tomatoes), as well as all other vegetables not falling into one of these categories. This last group is designated "garden vegetables" and includes, for example, green beans, corn, and broccoli. "Fruit" refers to whole fruit and fruit juice. In this article, fruit juices include citrus and other natural fruit juices, but exclude both fortified and non-fortified fruit "drinks." A list is available from the authors on request.

We report mean nutrient intake by number of servings (0, 1, 2, 3 or more) of fruit cross-classified by number of servings of vegetables. Nutrient intake was calculated by the National Center for Health Statistics, using USDA nutrient data, updated 1980. Because consumption of fruit and vegetables is associated with age, sex, and race,[1] we controlled for these factors in each cell of the cross-classification tables using a method equivalent to direct standardization. This method for finding means and the corresponding variance estimates, adjusted for age, race, and sex, is given in the Appendix. The method uses regression coefficients to obtain estimates of means.

Results

Because the NHANES II data were collected before the recommendations above were issued, they do not reflect their impact; rather the guidelines are used as a standard against which to describe the diet. In the spirit of both the USDA/DHHS and NAS guidelines, we define three or more servings of vegetables and two or more servings of fruit as meeting the vegetable and fruit guidelines, respectively. The proportions meeting these guidelines are shockingly small (Table 1). On the recall day, 27 percent met the vegetable guidelines and 29 percent met the fruit guidelines, while only 9 percent met both. In general, proportions were highest among older Whites but, even in this group, fewer than one-third ate the recommended servings of vegetables, and fewer than half ate the recommended servings of fruit. More men had adequate numbers of servings of vegetables than did women; the opposite was true for fruit. About two-thirds as many Blacks as Whites met these guidelines. Only about 5 percent of Black women and 7 percent of Black men consumed the recommended numbers of both fruit and vegetable servings.

Eleven percent ate neither fruit nor vegetables on the recall day, 45 percent had no servings of fruit, and 22 percent had no servings of a vegetable (Table 2). In contrast to the five or more servings of fruit and vegetables recommended, zero or one serving might be considered clearly inadequate. Thirty-five percent of Black adults and 27 percent of White adults had at most one serving of a food in the fruit/vegetable food group (data available from author on request).

Mean numbers of servings of fruit (1.08, SE = .03) and vegetables (1.77, SE = .02) were far below recommended levels and were lower for Blacks than for Whites (data available from author on request). Mean vegetable intake was

TABLE 1—Proportions (standard error) of Persons Ages 19–74 Who Met or Exceeded the USDA Guidelines, by Race, Sex, and Age Group: Estimates Based on Data from NHANES II, 1976–80.

	Age Group (years)			
	19–29	30–54	55–74	All Ages
Fruit[*] and Vegetables[†] Guidelines (2+ Fruits and 3+ Vegetables)				
White Males	.08 (.01)	.09 (.01)	.12 (.01)	.09 (.01)
Black Males	.05 (.01)	.08 (.02)	.08 (.02)	.07 (.01)
White Females	.05 (.01)	.08 (.01)	.15 (.01)	.09 (<.01)
Black Females	.04 (.02)	.03 (.01)	.08 (.02)	.05 (.01)
All	.06 (.01)	.08 (.01)	.13 (.01)	.09 (<.01)
Vegetable Guidelines (3 or more servings)				
White Males	.32 (.01)	.31 (.01)	.31 (.01)	.31 (.01)
Black Males	.22 (.04)	.20 (.04)	.23 (.05)	.21 (.02)
White Females	.21 (.01)	.25 (.01)	.28 (.01)	.25 (.01)
Black Females	.15 (.03)	.18 (.04)	.20 (.03)	.18 (.02)
All	.25 (.01)	.27 (.01)	.28 (.01)	.27 (.01)
Fruit Guidelines (2 or more servings)				
White Males	.21 (.02)	.25 (.01)	.35 (.01)	.26 (.01)
Black Males	.16 (.03)	.24 (.03)	.22 (.03)	.21 (.02)
White Females	.24 (.02)	.28 (.01)	.47 (.02)	.32 (.01)
Black Females	.24 (.05)	.20 (.03)	.32 (.05)	.24 (.02)
All	.22 (.01)	.26 (<.01)	.40 (.01)	.29 (<.01)

[*] Whole fruit and fruit juice

[†] Potatoes, salad, dried peas and beans, all other vegetables (denoted "garden": see text)

TABLE 2—Proportions (standard error) of Persons Ages 19–74 and Numbers in Sample, by Numbers of Servings of Fruit and/or Vegetables Consumed: Estimates Based on Data from NHANES II, 1976–80

Servings of Fruit*	Servings of Vegetables†				
	0	1	2	3+	All
0	.11 (<.01) 1,306	.11 (<.01) 1,244	.11 (<.01) 1,233	.11 (<.01) 1,247	.45 (.01) 5,030
1	.05 (<.01) 633	.07 (<.01) 814	.07 (<.01) 804	.07 (<.01) 847	.26 (<.01) 3,098
2	.03 (<.01) 325	.04 (<.01) 450	.04 (<.01) 486	.05 (<.01) 562	.15 (<.01) 1,823
3+	.02 (<.01) 261	.03 (<.01) 403	.04 (<.01) 481	.04 (<.01) 552	.14 (.01) 1,697
All	.22 (.01) 2,525	.26 (.01) 2,911	.26 (.01) 3,004	.27 (.01) 3,208	1.000 11,648

*Whole fruit and fruit juice

† Potatoes, salad, dried peas and beans, all other vegetables (denoted "garden": see text)

somewhat higher for males than females; the opposite was true for fruit and juice. The mean number of servings of garden vegetables, a group including the green and yellow vegetables emphasized by the NAS guidelines, was only 0.65 (SE = .01).

At least one serving of garden vegetables was eaten by 43 percent of the adult population, and there were no notable differences by race or sex (Table 3). Potatoes and salad were more popular among Whites than Blacks. Proportions eating fried potatoes declined with age, while those consuming non-fried potatoes increased (data available on request to author). Dried peas and beans, good sources of fiber, were eaten by only about 10 percent of the population.

About half the population consumed no servings or only one serving of a vegetable (Figure 1). Among those who reported only one serving of a vegetable, salad was the most popular, reported by 39 percent. Twenty-nine percent consumed a serving of potatoes as their only vegetable, and over half of these servings of potatoes were fried. Among those reporting two servings of a vegetable, some did not achieve variety: 9 percent had two servings of potatoes, 7 percent reported two servings of salad, and 5 percent ate two servings of dried peas and beans.

Only slightly more than half of the population (55 percent) had at least one serving of fruit or fruit juice. Of these, 27 percent had juice only, and 45 percent had whole fruit only (data available on request to author).

Table 4 shows average intake of dietary fiber, vitamin A, and vitamin C among persons consuming various numbers of servings of fruit and vegetables. These nutrient levels represent dietary intake from all food sources, but exclude vitamin supplements. Persons who consumed neither fruit nor vegetables had, on average, only 5.6 grams of fiber in their diet on the recall day, or about one-fourth of the 20–30 grams recommended by the National Cancer Institute.[13] The amount of fiber in these self-selected diets increased with the total number of servings of fruit and vegetables consumed, and this increase was similar for both fruit and vegetables. A total of five servings of a combination of fruit and vegetables was associated with an intake of approximately 17 grams of fiber. The recommended level of fiber intake was achieved only by those who had three or more servings of *both* fruit and vegetables, a group representing only 4 percent of the US adult population.

In the United States, it is estimated that the usual foods available to consumers provide about half of the total vitamin A activity in the diet as provitamin A carotenoids, found in plant products, and the other half as retinol.[14] Among those eating both fruit and vegetables, the US Recommended Daily Allowance (RDA) of 5000 IUs of vitamin A was attained only by those having at least two servings, while five servings of a combination of fruit and vegetables were associated with mean vitamin A intake of about 8000 IUs. If either vegetables or fruit were eaten to the exclusion of the other (true of over half the population), the US RDA was reached only among those having three or more servings.

For vitamin C, fruit and vegetables represent the primary source in the US diet,[10] and Table 4 shows that fruit in particular is associated with higher intake. Among those consuming no fruit, the US RDA of 60mg was attained only by those who had three or more servings of vegetables. This is notable in view of the fact that 45 percent of the US population consumed no servings of fruit on the day of the survey.

Table 5 shows caloric and fat intake by servings of fruit and vegetables. Caloric intake increased as numbers of servings of either food increased. However, the additional calories were not due to the fruit or vegetables alone, as fat intake also increased substantially. This is presumably due in part to fats such as butter and margarine added to potatoes and other vegetables and to oils in salad dressings; people who eat more servings of fruit and vegetables may also eat more food in general. The increase in dietary fat intake associated with increased servings of vegetables was greater than that associated with increased servings of fruit (data available on request to author). Those who ate no fruit and three or more servings of vegetables had an average of 95 grams of fat in their diets, while those who ate three servings of fruit and no vegetables had 78 grams.

TABLE 3—Proportions (standard error) of Persons Consuming One or More Servings of a Vegetable, by Race, Sex, and Type of Vegetable: Estimates Based on Data from NHANES II, 1976–80

Type of Vegetable	Whites		Blacks		All
	Males	Females	Males	Females	
Garden Vegetables*	.42 (.01)	.44 (.01)	.41 (.03)	.44 (.02)	.43 (.01)
Potatoes					
Fried	.22 (.01)	.13 (.01)	.15 (.02)	.10 (.01)	.17 (<.01)
Not Fried	.28 (.01)	.26 (.01)	.20 (.02)	.20 (.02)	.26 (.01)
All	.46 (.01)	.37 (.01)	.33 (.03)	.30 (.03)	.40 (.01)
Salad	.39 (.01)	.40 (.01)	.19 (.02)	.24 (.02)	.37 (.01)
Dried Peas and Beans	.11 (.01)	.08 (.01)	.13 (.02)	.10 (.02)	.10 (.01)

*All vegetables except potatoes, salad, and dried peas and beans. Examples are green beans, carrots, and corn.

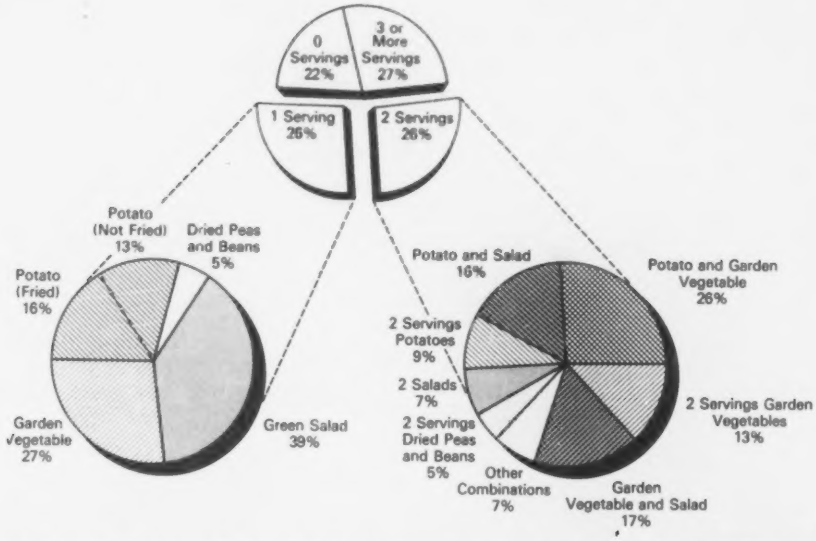

FIGURE 1—The top pie chart shows the percentages of individuals who consumed zero, one, two, or three or more servings of vegetables on the recall day. Percentages are estimates based on data from NHANES II, 1976–80. The lower left pie chart displays the type of vegetable consumed by those who consumed only one serving of a vegetable. The pie chart on the lower right displays the types of vegetables consumed by those who had exactly two servings of a vegetable. The category "garden vegetables" includes all vegetables except potatoes, salad, dried peas and beans. Examples are green beans, carrots, and corn.

When we express fat as a percent of calories, there was a small increase with increasing vegetable intake, approximately two percentage points from the lowest to the highest serving category. In contrast, with increasing fruit consumption, the percent of calories from fat fell approximately five percentage points from the lowest to the highest consumption category. Percent of calories from saturated fat behaved similarly, reaching its highest value among those who consumed no fruit and two or more vegetables, and its lowest value among those who consumed three or more servings of fruit and no vegetables (data available from authors on request).

Although the percent of calories from fat increased slightly with increasing vegetable consumption, this was associated with an increase in the ratio of polyunsaturated fats to saturated fats (P/S ratio), the direction recommended by the American Heart Association[15] (data available from authors on request). This may reflect the use of unsaturated oil in salad dressings.

Despite the increase in caloric intake and dietary fat with increasing numbers of servings of fruit and vegetables, there was no associated increase in obesity (data available from authors on request). A commonly used measure of obesity, the Quetelet Index (weight(kg)/height(m)^2) was relatively constant across all serving categories, ranging from 24.9 (SE = .27) to 25.6 (SE = .18). This may reflect differences in energy expenditure associated with higher caloric levels.

Discussion

"In this land of plenty millions of Americans aren't eating wisely. Not because they haven't enough to eat, but because they eat too many of the wrong things or too little of the right,"[17] according to the USDA. Based on 24-hour recall data, we found that 91 percent of the adult US population did not meet the USDA/DHHS recommendation that two to

TABLE 4—Mean Intake of Various Nutrients (Standard Errors) by Number of Servings of Fruit and Vegetables, Adjusted for Age, Race, and Sex: Estimates based on Data from NHANES II, 1976–80

	Servings of Vegetables†			
	0	1	2	3 or more
*Servings of Fruit**	Dietary Fiber (g)			
0	5.6	6.62	9.7	13.8
	(.2)	(.2)	(.3)	(.5)
1	7.7	8.8	10.5	14.8
	(.2)	(.3)	(.3)	(.4)
2	9.9	11.0	13.2	16.6
	(.3)	(.3)	(.3)	(.3)
3 or more	13.8	14.7	17.2	21.0
	(.5)	(.3)	(.4)	(.5)
	Vitamin A (IU)			
0	2998	3333	4880	5882
	(167)	(189)	(430)	(269)
1	3804	5364	5311	6789
	(314)	(649)	(279)	(325)
2	3876	5372	5855	7934
	(253)	(451)	(283)	(456)
3 or more	5330	7041	8924	10393
	(393)	(403)	(615)	(526)
	Vitamin C (mg)			
0	25.9	38.3	51.4	79.6
	(2.1)	(1.7)	(1.8)	(3.4)
1	73.0	95.5	102.7	130.0
	(2.0)	(4.4)	(2.7)	(3.7)
2	114.4	123.5	145.0	163.0
	(6.9)	(3.3)	(4.6)	(3.9)
3 or more	218.0	211.8	222.8	248.3
	(12.0)	(10.7)	(7.0)	(9.7)

USRDA for vitamins A and C is 5000 IU and 60 mg respectively. The National Cancer Institute recommends a daily fiber intake of 20–30g with an upper limit of 35 g.

*Whole fruit and fruit juice

†Potatoes, salad, dried peas and beans, all other vegetables (denoted "garden": see text)

TABLE 5—Mean Caloric and Fat Intake (standard errors) by Number of Servings of Fruit and/or Vegetables, Adjusted for Age, Race, and Sex (Estimates based on Data from NHANES II, 1976–80)

	Servings of Vegetables[†]			
	0	1	2	3 or more
Servings of Fruit[*]	Calories			
0	1633	1817	1950	2230
	(30)	(26)	(33)	(47)
1	1741	1890	1941	2264
	(34)	(36)	(32)	(70)
2	1924	1937	2029	2249
	(51)	(48)	(43)	(49)
3 or more	2112	2137	2173	2369
	(84)	(59)	(57)	(50)
	Fat *(g)*			
0	64.4	75.6	84.4	95.4
	(1.3)	(1.8)	(1.8)	(2.1)
1	69.1	79.7	81.2	98.3
	(1.9)	(1.9)	(1.6)	(3.5)
2	75.8	77.2	81.3	93.5
	(2.9)	(2.5)	(2.3)	(3.3)
3 or more	77.6	83.2	83.3	92.0
	(3.7)	(3.1)	(3.0)	(2.4)
	Percent Calories from Fat[‡]			
0	37.6	40.1	40.6	40.3
	(.3)	(.4)	(.3)	(.3)
1	36.1	38.9	39.6	39.8
	(.3)	(.4)	(.3)	(.4)
2	36.1	36.5	36.9	37.8
	(.5)	(.5)	(.5)	(.3)
3 or more	32.7	35.0	[illegible]	34.8
	(.7)	(.6)	(.5)	(.4)

[*] Whole fruit and fruit juice

[†] Potatoes, salad, dried peas and beans, all other vegetables (denoted "garden": see text)

[‡] Calculated as percent of nonalcohol calories.[16]

three servings of fruit and three to five servings of vegetables be eaten daily.

Our method of defining a serving in terms of the number of grams eaten at any given meal, as opposed to the total number of grams consumed during the entire 24-hour period, may have resulted in some bias in the estimation of number of servings. Small portions (weighing less than an ounce) might sum to a serving over the course of a day, resulting in an underestimation. This bias is probably minimal because of the small lower limits (one ounce of fruit or vegetable) used in the estimation. Conversely, these small limits might result in an overestimation of the number of servings; our lower limit is equivalent to less than half of a serving as defined by the USDA. The second bias is likely larger than the first and, if so, the results presented in this article are unduly optimistic.

While a single day's food record cannot be used to characterize an individual's usual diet, 24-hour recall data can provide estimates of group means.[7] Our finding that 11 percent had no servings of either fruit or vegetables on the recall day is in close agreement with data from the NHANES II food frequency questionnaire in which 12 percent reported that they eat fruit and vegetables less frequently than daily, and with 24-hour recall data from the Nationwide Food Consumption Survey (NFCS), conducted by the USDA in 1977–78, in which 10 percent of women and 8 percent of men ages 19–50 reported eating no fruit or vegetables on the recall day.[18]

A comparison of 24-hour dietary recall data among adults ages 19–50 from the NFCS and from the 1985 and 1986 Continuing Survey of Food Intakes (CSFII),[19–21] also conducted by the USDA, suggests that there has been little change in the percentages using these foods. The proportion of men reporting eating any vegetables on the recall day decreased from 89 to 85 percent between 1977 and 1985, while the proportion reporting fruit consumption was virtually unchanged: 44 percent in 1977 and 43 percent in 1985. For women, changes were also minimal: in 1977, on the recall day, 84 percent reported eating a vegetable, compared to 83 percent in 1985; 50 percent reported eating fruit in 1977 and 47 percent in 1985. CSFII data for women in 1986 were similar. A study of food consumption trends based on NFCS and CSFII data reported that proportions of women using green and yellow vegetables and "lower-fiber" vegetables such as lettuce and tomatoes increased slightly during this period, while proportions using "higher-fiber" vegetables such as corn, green peas, and cooked cabbage decreased.[22] While data from the USDA surveys are directly comparable, these surveys differ somewhat from the NHANES II survey in sampling and in data collection methods. Further, the USDA vegetable group is slightly more comprehensive than ours, and their fruit group excludes some citrus fruit juices, such as lemonade, that were included in our study. Recent promotion of dietary fiber may have produced some increase in the consumption of fruit and vegetables, but is unlikely to have resulted in changes of the magnitude needed to meet the guidelines.

Results of a small survey (298 adults selected by random digit dialing) conducted in two counties in California in 1987[23] were similar to our findings. Fifty-two percent reported that they had eaten one or more fresh fruits on the previous day*; we found that 39 percent ate whole fruit on the recall day. In the California survey, 42 percent reported eating green salad and 67 percent reported eating one or more vegetables on the recall day. In our data, 37 percent had a salad and 43 percent had at least one serving of a garden vegetable on the recall day.

Dietary guidelines have emphasized the importance of a balanced and varied diet. Diets that include either no servings or very few servings of fruit and vegetables lack both balance and variety. We found that almost 50 percent of the population had at most a single serving of a vegetable; even among those who had two servings of a vegetable on the recall day, 21 percent had two servings of the same vegetable (other than a garden vegetable). For fruit and fruit juice, about three-fourths had at most a single serving.

Studies that have assessed health benefits from vegetable consumption found them primarily for garden vegetables.[24-26] We found that less than half the population had a garden vegetable on the recall day.

As in the paper on food choices,[1] we found that consumption of fruit and vegetables (with the exception of garden vegetables) was lower among Blacks than Whites, even when income was taken into account (data available on request to authors). Not only did fewer Blacks than Whites meet the guidelines, but larger proportions had at most a single serving on the recall day.

Consumption of the recommended five servings of fruit and vegetables is consistent with a nutritionally adequate diet as well as with the health recommendations cited above. For example, diets including multiple servings of these foods provide substantial amounts of dietary fiber and vitamins A and C. While caloric intake increases with the numbers of servings of both fruit and vegetables, this increase is not associated with greater obesity. As more servings of vegetables are eaten, more grams of fat are consumed, but the P/S ratio also rises.

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*Data provided by S. B. Foerster, personal communication.

The fact that fruit and vegetables are important sources of dietary fiber may not be well understood by the general public. In our data, a diet that included five servings of fruit and vegetables provided approximately 17 grams of fiber; at this level, the addition of two slices of whole grain bread would bring the total to 20 or more grams. Nutrition education campaigns might well emphasize that consumption of the recommended five-a-day is important in achieving the dietary fiber goals, as well as in increasing micronutrient intake.

Evidence has been accumulating on the relation of fruit and vegetable consumption to health, as is reflected in the NAS recommendations to increase consumption of these foods. Numerous studies suggest that we can help prevent several types of cancer by increasing our fruit and vegetable consumption.[24-26] However, the discrepancy between the America diet and the fruit and vegetable guidelines, as shown in a one day recall, is so enormous as to call into question the societal commitment to these guidelines. The public is either unaware of the likely benefits of including these foods in the daily diet or, although aware, is unwilling or cannot afford to follow the guidelines.

Fruit and vegetable consumption has never been the object of a national campaign. A recent editorial in this Journal called for "the most vigorous, collaborative nutrition campaign ever launched"[27] to make the public aware of the connection between food intake and health outcome. Any such campaign should take into account the fruits and vegetables most frequently eaten, encouraging increased consumption of those already relatively popular, and targeting demographic groups where consumption is particularly low. In addition to public education, societal commitment might include actions to improve the affordability of many fruits and vegetables and to increase the numbers of servings obtainable through school lunches, meals on wheels, and other food programs, as well as industry actions to promote increased consumption of these foods.

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APPENDIX

We define K indicator (dummy) variables for the K + 1 cells in our classification (K = 15 for the 4 x 4 cross-classification of numbers of servings of fruits and vegetables.) The remaining P - K variables, $X_{K+1}, \ldots, X_p$, represent indicator variables for the categories into which the adjustment variables have been divided (in our case, the three age, two race, and two sex categories generate four additional dummy variables). The continuous variable for which we are adjusting is denoted Y_k, $k = 1, \ldots, K$ (for example, Y_k might be number of calories consumed by those in the k^th cell). Then the adjusted mean for the k^th cell, $k = 1, \ldots, K$, is given by the following regression model:

$$\bar{Y}_k = b_0 + b_k + b_{K+1}\bar{X}_{K+1} + \ldots + b_P\bar{X}_P$$

and the adjusted mean for the $(K + 1)^{th}$ cell is given by

$$\bar{Y}_{K+1} = b_0 + b_{K+1}\bar{X}_{K+1} + \ldots + b_P\bar{X}_P$$

where $\bar{X}_{K+1} \ldots \bar{X}_P$ are means for the whole sample, or, in our case, simply the proportion (in the population) in each age, sex, or race category. A simple interpretation can be given to the adjusted means. The adjusted mean for the k^{th} cell is the mean that would have been generated by the data if the population in the k^{th} cell had the same distribution (with respect to the race, sex, and age categories) as the entire population.

The estimated variance of the adjusted mean can be found by using the estimated variances and covariances of the b's, obtained from the SURREGR procedure,[8] software that takes into account the complex sampling design and the sampling weights.

Executive Women and Health: Perceptions and Practices

JUDITH H. LAROSA, PHD, RN

Abstract: The purpose of this study was to obtain a socioeconomic/health profile of a select group of executive women, to understand more about their personal and professional lives, and to examine how these factors relate to their overall health. The data were obtained from a self-administered 73-item questionnaire that was mailed during spring 1987 to the 1,000 members of a professional executive women's organization with 15 chapters across the United States. Findings suggest that the women in executive positions do not necessarily compromise their health. In comparison with a group of age/gender/education matched working women, the overall wellness and risk assessment scores were remarkably similar. The study group, however, reported greater life satisfaction, stronger social support, and excellent health status; the overwhelming majority was satisfied with their personal and professional lives and believed that they were in control of both. This perception, coupled with the relatively high wellness scores, suggests that on average this group of executives may be in better health than had been predicted as women rose to executive positions within organizations. (*Am J Public Health* 1990; 80:1450–1454.)

Introduction

The role of women has changed—both inside and outside the home. Since the beginning of the twentieth century, many women have moved from the traditional rural culture of child-rearing and farm work to the modern urban culture of nannies and executive boardrooms.

Although the move from farm to city and from home to office has been evolving since the early 1900s, it was during World War II that the major and more permanent change occurred. In 1890, women constituted 18 percent of labor force participation, by 1945 that number had risen to 36 percent, and in 1989, women represented 54 percent of the labor force participation.[1,2]

During World War II, as throughout history, women took on the work responsibilities vacated by men who left home and jobs to fight wars. Unlike the aftermath of previous wars, however, American women did not follow historical precedent and return home to their traditional responsibilities. They continued to work outside the home and over time began to take on occupations previously held by men—medicine, the law, engineering, politics, and others. Gradually they began to take on the power and responsibilities associated with these occupations. In many cases, women added these outside responsibilities to the ones they already had at home—managing the children and the household.

Yet, little attention has been focused on the impact of these changes. Studies are only just now emerging that begin to examine the complex nature of the home/work/health relationships on working women. Some investigators have suggested that functioning in different roles poses health problems for working women,[3–5] while others have suggested that it is beneficial.[6–8] Some have suggested that a critical factor in these relationships is the perception that the woman holds about her life situation.

Prior health studies focused primarily on women in blue-collar occupations, or on women in white-collar occupations who worked at relatively low to mid-level positions.[3–24] Existing studies of executive women usually do not include health data.[25,26] Indeed, only three studies were identified that focused on the health of women in executive positions—either in very large companies or in the federal government.[27–29] The purpose of this study, therefore, was to obtain a socioeconomic/health profile of a select group of executive women, to understand more about their personal and professional lives, and to examine how these factors relate to their overall health.

Methods

The data were obtained from a self-administered questionnaire that was mailed during spring 1987 to the 1,000 members of a professional executive women's organization with 15 chapters across the United States. Entry into the organization is by invitation only and is based on criteria that insists on each member having "substantial influence in her field . . . regardless of title." The organization criteria also states that "there are many influential and powerful women outside large organizations, thus the size of the organization should not be considered as an important criterion." Occupations include chief executive officers, lawyers and judges, physicians and scientists, administrators, politicians, the clergy, artists, and writers. The majority, 54.8 percent, listed their title as executive. The mean income of the respondents was $133,000 per annum; median income was $96,000. The income distribution was positively skewed. The income levels ranged from no income for a Roman catholic nun who was a hospital administrator to several chief executive officers who reported over $3 million in salary per annum. Because of the wide range of incomes, further analyses of income against other key variables are underway.

The questionnaire consisted of 73 items in four major categories—sociodemographic, general health, job satisfaction, and home/work/health relationships. Sociodemographic items were based on those used in the National Health and Nutrition Examination Surveys and National Health Interview Surveys.[30,31] Participants were also asked job-related questions and queried about who bears the major responsibility for making major financial decisions and arranging child care.

The general health section consisted of the 1984 32-item[*] Centers for Disease Control Health Risk Appraisal (CDC HRA) with the exception of one item reflecting exposure to

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Address reprint requests to Judith H. LaRosa, PhD, RN, Executive Women's Health Study, 1958 Dundee Road, Rockville, MD 20850. Dr. LaRosa is also with the Department of Health Education, University of Maryland, College Park. This paper, submitted to the Journal February 20, 1990, was revised and accepted for publication April 6, 1990.

[*] One question was dropped from the original questionnaire at the request of the organization's board. The question was considered offensive and board members felt that such a question would seriously impair the willingness of members to respond. The excluded question was "How many of the following do you usually do? Hitch-hike or pick up hitch-hikers? Carry a gun or knife for protection? Keep a gun at home for protection? Criticize or argue with strangers? Live or work in a high crime area? Seek entertainment at night in high-crime areas or bars?

violence.[32] This instrument was used for two reasons. In addition to providing an assessment of the individual's perception of her health status, the CDC HRA also provided a limited assessment of the respondent's level of personal satisfaction—an important psychosocial variable previously shown to be related to the development of disease.[9–11]

The job satisfaction section consisted of two scales developed by Karasek and colleagues—a five-item job satisfaction scale and a nine-item decision latitude scale.[33–37] These variables were added because professional satisfaction and perception of job mastery are believed to be related to health.[4,9–11,33–37]

Perceptions about control over their lives and the impact of their personal and professional lives on physical health and mental well-being were examined using five questions developed by the investigator. Finally, two open-ended questions were included at the end of the questionnaire for the respondents to add comments on their personal and professional lives and their health, and to reflect on the study. Analyses of the responses to these questions are not included in this paper; they will be the subject of a future analysis.

After a three-phase mailing, designed to maximize the response rate, 66 percent (660/1,000) of the members had responded. Of those 660, 590 cases (59 percent of the total) were complete and usable; 545 of the cases represented executives who were employed full-time. They are the focus of this paper.

Complete data for comparing respondents and nonrespondents were available on only two factors: education and marital status. No differences were found between the two groups on these factors. While differences may have existed, other demographic factors such as age, number of years married, spouse education and income, level of personal and professional satisfaction, and the entry criteria for membership in the organization suggest that this particular population of executive women is fairly homogeneous. The assumption that the respondents and nonrespondents are similar seems reasonable.

Results

The 545 respondents were predominantly White, middle-aged, well-educated, well-paid executives. The majority were married and had an average of two children. Almost 19 percent of the children were 18 years of age or younger. The majority of the spouses, like their wives, were well-educated executives. Table 1 summarizes the selected demographic findings and job characteristics.

Although the majority of the couples had equally well-paying, executive positions, the distribution of jobs at home was not as equitable. While the responsibility for making major financial decisions was a shared responsibility between the partners, the responsibility for arranging child care usually fell to the woman. Seventy-three percent of the respondents reported arranging child care 75 percent or more of the time.

Over half the respondents have had at least two job changes in the past 10 years, usually representing an advancement. While the majority of the respondents worked for a firm other than their own, 28 percent owned their own company. The majority reported high levels of job satisfaction (median equaled 90 on a scale of 0 to 100) and decision latitude (median equaled 86 on a scale of 0 to 100)—indicating an ability to make autonomous decisions regarding their work direction and pace.

TABLE 1—Profile of 545 Full-Time Women Executives

Factors		
Age (years) (Mean/Median)	48.7/48.0 years	
Race (% White)	90.1	
Education (% ≥ Masters)	57.0	
Spouses' Education (% ≥ Masters)	66.1	
Married (%)	59.5	
Married (years)	20.0	
Children (% with)	73.0	
Children (Average Number)	2.0	
Children (% with ≤ 18 years of age)	19.0	
Annual Personal Income (Mean/Median)	$133,785/$96,000	
Spouse's Annual Personal Income (Mean/Median)	$143,102/$90,000	
Annual Total Family Income (Mean/Median)	$271,401.00/$175,000.00	
Hours worked per week (Mean/Median)	54.0/54.0	
Self-employed (% who owned firm)	27.5	
Occupation	Respondent	Spouse
Business Executive	54.8	45.7
Lawyer	11.2	17.2
Physician	0.0	6.2
Government Official	9.0	0.0
College Administrator/Professor	6.2	4.2
Other	18.8	26.0

NOTE: Other includes editor, publisher, author/correspondent, physician, CPA, architect, psychologist, performer/artist

Over 56 percent of the respondents believed that they had a great deal of control over their personal and professional life. An additional 31 and 34 percent in each category (personal and professional, respectively) believed they had a moderate amount of control. In addition, the majority of respondents (77 percent) reported being at least mostly satisfied with their lives.

Respondents furthermore reported that their personal and professional lives had a strongly positive effect on both physical health and mental well-being. In both personal and professional categories, however, the positive effect on mental well-being was stronger than the positive effect on physical health (data available from author upon request).

In general, respondents believed that they were in good health and reported maintaining positive health practices. The mean CDC HRA wellness score, based on a 100-point scale, was 86, with a median score of 87.

This score requires a reference point to indicate where this group falls on the 100-point continuum in comparison with other groups. Although no entirely appropriate comparison group who had completed the CDC HRA could be found for these senior-level executive women, their responses were compared with those from a matched group of working women in the Fitness Research Center data base at the University of Michigan.** Each of the participants in the study group (N = 533)*** was matched on age and education with 20 working women from the Fitness Research Center data base for a total of 10,660. Because neither of the two groups is necessarily representative of all working women nor all executive women, the findings should be interpreted with caution.

**The Fitness Research Center at the University of Michigan maintains a current data base of over 100,000 health risk appraisals collected from the University of Michigan Business School and from diverse data bases provided by investigators from across the United States.

***Twelve cases of the original 545 were not used either because of insufficient data or because respondents were over 70 years of age. Thus, the 533 used represented complete cases.

Table 2 summarizes the CDC HRA health findings from the study reported here and compares them with the CDC HRA health findings from the matched reference group.

The average wellness scores and the risk assessment scores between the two groups were remarkably similar. Yet, further examination reveals several important differences. Fewer of the study group were married than were the matched group with the major difference occurring among those who were divorced. Fewer of the study group smoked than did the matched group, yet far more of the study group reported drinking than did the matched group. Although the number who reported heavy drinking (≥25 drinks of alcohol per week) was small, the study group had over three times as many heavy drinkers as the matched group.

Blood pressure and total blood cholesterol levels were reported as slightly lower among the study group than the matched group.† However, this should be interpreted carefully as the study group of executives was far less knowledgeable about their blood pressure and blood cholesterol values than the matched group. The majority of the executives in the study group did not know either value.

Differences existed between the two groups concerning regular exercise as well as miles driven and seat belt use. The study group exercised more regularly (49 percent) than did the matched group (44 percent). The executives drove, on average, considerably more miles per year than did the matched group and were far more likely to use seat belts (80 percent) than did the matched group (36 percent).

Discussion

This study provides some preliminary data that suggest that the women in executive positions do not necessarily compromise their health. Indeed, this particular group of executive women perceived themselves to be quite healthy, which was supported by their overall wellness and risk assessment scores. In comparisons with a group of age/gender/education-matched working women, the overall wellness and risk assessment scores were remarkably similar. The study group reported greater life satisfaction, stronger social support, and excellent health status. The overwhelming majority of the study group were satisfied with their personal and professional lives and believed that they were in control of both. This perception, coupled with the relatively high wellness scores, suggests that this group of executives may be in better health than had been anticipated as women rose to executive positions within organizations.

The greater numbers of those who drank and who drank heavily, as well as the greater number of divorcees among the study group, requires further examination. Perhaps among a sub-set of these women, there are those for whom the executive lifestyle had a negative effect.

That the executives drink more is not entirely unexpected. Although fewer business lunches and dinners today are conducted with alcohol as an important social ingredient, alcohol is still consumed at such meetings. It may be that alcohol consumption is perceived as a necessary social part of doing business. Or, it may be that these executives really do consume alcohol more often than do other types of working women, possibly as a way of reducing stress. Yet, other behaviors that have been linked to stress reduction,

TABLE 2—Demographics, Disease Risk Factors, and Health Practices and Perceptions among the Study Group and the Matched Group

Factors	Study Group (N = 533*)	Matched Group (N = 10,660)
Sociodemographics		
Age (Mean)	49	49
Marital Status (%)		
Married	58.5	65.3
Single	12.2	10.5
Divorced	20.1	14.1
Separated	2.6	2.2
Widowed	5.3	7.4
Race/Ethnic Group (%)		
White	89.9	89.9
Black	7.1	5.2
Hispanic	1.9	1.2
Asian	0.6	2.5
Native American	0.0	0.9
Other	0.6	0.2
Education (%)		
≤ High School	7.1	7.1
Some College	2.1	2.1
College or Professional	90.6	90.6
Risk Factors		
Height/Weight (Mean)		
Height (Inches)	65.5	65.5
Weight (Pounds)	136.7	137.4
Smoking (%)		
Smoker	12.0	17.7
Ex-smoker	38.3	23.3
Never Smoker	49.7	58.9
Among Smokers		
Average/day	11.4	19.0
Among Ex-smokers		
Years Stopped	5.4	10.3
Blood Pressure (%)		
High SBP** ≥ 140 mm Hg	5.5	21.9
High DBP** ≥ 90 mm Hg	0.4	17.2
Mean SBP (mm Hg)	117.8	120.1
Mean DBP (mm Hg)	75.0	77.1
Know Blood Pressure	44.5	68.5
Blood Cholesterol (%)		
BC ≥ 5.7 mmol/L	25.6	36.1
Mean Total BC (mmol/L)	5.0	5.6
Know Blood Cholesterol	16.1	25.0
Alcohol Use (%)		
Drinker	72.2	45.3
Ex-drinker	2.1	2.3
Never Drinker	25.7	52.5
Among Drinkers Only (%)		
Light (≤ 7/week)	57.4	72.5
Medium (8–24/week)	39.2	26.9
Heavy (≥ 25/week)	3.1	0.6
Average/week	8.1	7.8
Drug Use		
Non (Rarely) Use	92.3	93.5
Light (Occasional)	5.6	4.8
Heavy (Almost daily)	2.1	1.6
Health Practices		
Physical Activity		
Regular (≥ 3 times/week)	49.3	44.1
Occasional	37.1	43.6
Little	13.5	12.3
Sleep		
≤ 6 hours/night	24.6	27.1
7–8 hours/night	72.6	72.0
≥ 9 hours/night	2.8	1.7
Pap Smear		
Annually	68.3	63.5
Every 3 Years	22.0	22.9
< Every 3 Years	7.1	9.7
Almost Never	0.2	0.7
Breast Self-examination		
Monthly	21.2	36.7
Every Few Months	40.3	35.5
Rarely/Never	38.5	27.8
Annual Rectal Examination	47.1	37.1

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†Where blood pressure and blood cholesterol data are missing, the CDC HRA program automatically assigns an age/gender/race mean to that case.

TABLE 2—Continued

Factors	Study Group (N = 533*)	Matched Group (N = 10,660)
Seatbelt Use		
Regular (≥ 75%)	80.5	36.4
Semi-regular (<75%)	6.0	37.4
Non-user	13.5	26.1
Average Use (%)	82.9	47.4
Average driving miles/years	26,538	11,191
Perceptions		
Life Satisfaction		
Completely	77.3	60.4
Quite	21.0	28.5
Somewhat	1.3	9.9
Not Very	0.4	1.2
Perceived Social Support		
Strong	69.6	52.7
Average	21.2	40.5
Below Average	4.1	5.9
Unsure	2.1	1.1
Perceived Health Status		
Excellent	61.4	36.0
Good	34.0	55.8
Fair	3.9	7.6
Poor	0.8	0.8
Risk Assessment		
Average Risk Age (years)		
Actual	49.6	48.6
Appraised	48.1	47.4
Achievable	46.0	44.9
Projected Death Rate All Cause Death/1,000 Persons during Next 10 Years		
Actual	77.4	75.1
Appraised	68.3	71.6
Achievable	57.7	55.4
Wellness Score (Mean)	85.5	78.3

*533 represents cases which are complete.
**SBP = Systolic Blood Pressure; DBP = Diastolic Blood Pressure

such as smoking and overeating, appear to be less common with this group of executives compared to the matched group. In fact, more positive stress-reducing activities, such as regular exercise, appear to be practiced by almost half of the executive group.

It is not surprising that at the time of the study (1987) many of the participants and matched group were unaware of their blood cholesterol levels. The National Cholesterol Education Program had just been initiated and related public education was in its infancy. What is surprising is the number of executives who did not know their blood pressure values. The National High Blood Pressure Education Program and its public education efforts have been strong in the United States for the past 17 years. Knowledge about one's "number" and awareness about the disease has been increasing since the inception of this national educational effort. Indeed, knowledge about blood pressure and action to control it has been highly correlated with education and income, especially among females. The reason for this lack of knowledge about blood cholesterol and blood pressure levels is puzzling and is worthy of further investigation.

The greater effort expended on exercise is certainly within keeping among men with higher education. Now Matthews, *et al*, have reported similar findings among women with college and advanced degrees, e.g., the higher the education, the greater the energy expenditure.[38]

While this study may begin to dispel some of those notions of compromised health among women executives, it raises far more questions than it answers. Clearly the study has limitations. It captures only one moment in time and it does not represent all executive women. However, it provides suggestions that might direct further study. The study and the review of the literature argue vigorously for more investigation in this particular area of executive working women, and in the larger concern of the interrelationships of home, work, and health for all working women.

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Innovations in State and Local Government Awards Program Invites 1991 Applications

The Innovations in State and Local Government Program has issued a call for applications for its 1991 Innovations Awards Program to recognize creative government efforts at the state, county, municipal, special authority, and tribal levels. Ten winners, selected by a national committee of leaders from government, education, business, and the media, will receive grants of up to $100,000 each by the Ford Foundation.

Entries will be evaluated according to: the program's novelty; its significance in addressing important local and national problems; the value it brings to its clients; and the degree to which it can be replicated by other jurisdictions.

For more information and a 1991 application, contact: Doug Carpenter, Operations Manager, Innovations in State and Local Government, A. Alfred Taubman Center, 79 John F. Kennedy Street, Cambridge, MA 02138; (617) 495-0557.

To be considered, all applications must be received no later than 5:00 p.m., Wednesday, January 9, 1991.

Day Care Characteristics Associated with *Haemophilus influenzae* Disease

Jay D. Wenger, MD, Lee H. Harrison, MD, Allen Hightower, MS, Claire V. Broome, MD,
Haemophilus influenzae Study Group*

Abstract: To identify characteristics of day care facilities associated with *H. influenzae* disease, we compared 92 licensed facilities in which a case of *H. influenzae* disease had occurred with randomly selected facilities at which no cases occurred. Matched univariate analysis showed that personnel at facilities where *H. influenzae* disease occurred were more likely than those at control facilities to use towels or handkerchiefs to wipe children's noses, admit children who were not toilet trained or had diarrhea ("liberal fecal policy"), had a narrower age range, were more likely than control facilities to be for-profit and less likely to use volunteers. In a multivariate model that adjusted for age range, profit status and liberal fecal policy, towel or handkerchief use (OR 5.5, 95% CI: 1.1, 30) was the only variable independently associated with case facilities. This is the first association of a specific day care practice with *H. influenzae* disease. (*Am J Public Health* 1990; 80:1455–1458.)

Introduction

Whereas 20 years ago most children were cared for at home, today most women with children less than 6 years of age are working outside the home, and most of their children receive care at day care facilities.[1,2] Numerous studies have suggested that children attending day care centers were more likely than those cared for at home to suffer from a variety of infectious diseases.[3–9] In particular, attendance at day care facilities is a major risk factor for primary invasive *Haemophilus influenzae* disease.[10–12] Since nearly all *H. influenzae* disease is primary (i.e., is not a result of transmission from another known case), we performed a case-control study comparing licensed day care facilities where a case of *H. influenzae* disease occurred with licensed day care facilities which did not have cases of disease in the same period.

Methods

Surveillance for Invasive *H. influenzae* Disease

Laboratory-based active surveillance for invasive *H. influenzae* disease was performed in four states (Tennessee, Missouri, New Jersey, and Washington) and Los Angeles County, California from January 1, 1986 to December 31, 1986. The details of this system have been previously reported.[13] A contact in each acute care hospital in the area was responsible for reporting to the surveillance coordinator all cases of *H. influenzae* disease from that hospital. A child was considered to be a case of *H. influenzae* disease if *H. influenzae* was cultured from a normally sterile site. When the project ended, surveillance coordinators reviewed ICD-9 discharge diagnosis codes in each hospital to evaluate the sensitivity of the surveillance system.

Selection of Case Day Care Facilities

A child was considered a day care facility attendee if he/she was being cared for along with at least one other unrelated child for at least four hours during the week before becoming ill with *H. influenzae*. The surveillance coordinator contacted the parents of all children between the ages of 1 and 60 months by telephone to determine day care facility attendance. A day care facility was entered into the study only if it appeared on the list of licensed day care facilities provided by the state or county health departments, since no listing of unlicensed facilities existed to select controls for cases attending unlicensed facilities. The director of each case day care facility was contacted by telephone and, after explanation of the study, offered the opportunity to participate. If a case day care facility refused to participate, the case was eliminated from the study. If the director of the case day care facility agreed to cooperate, the surveillance coordinator obtained the case center's characteristics via questionnaire at that time.

Selection of Control Day Care Facilities

For each enrolled case day care facility a size-weighted, randomly ordered list of licensed day care facilities in the surveillance area was generated. The first day care facility on the list was chosen as a control day care facility if it:

- matched the case day care facility for type, i.e. day care home (in a residence) or nonresidential day care center;
- was not a case day care facility itself; and
- had not already been selected previously as a control.

The director of each potential control day care facility was contacted by telephone until a day care facility was found that would have cared for children in an age range including the age of the case at the time of case admission and was willing to participate.

Statistical Methods

Univariate and multivariate models were evaluated using matched conditional logistic regression.[14] Analyses performed on subsets of the data (for example, for cases under two years of age or for day care homes vs day care centers) contained only sets where both the case and control met the

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From the Meningitis and Special Pathogens Branch and Statistical Services Activity, Center for Infectious Diseases, CDC; and the *Haemophilus influenzae* Study Group. Address reprint requests to Jay D. Wenger, MD, Building 1, Room 5409, CO9, Centers for Disease Control, Atlanta, GA 30333. This paper, submitted to the Journal December 11, 1989, was revised and accepted for publication May 22, 1990.

*The *Haemophilus influenzae* Study Group: Sandra L. Sitze and Margaret Spurrier (Missouri Department of Health, Jefferson City, MO); Robert C. McCready, David R. Cundiff, and Maureen C. Farrell (New Jersey Department of Health, Trenton, NJ); Janice D. Harwell (Washington State Department of Social and Health Services, Seattle, WA); Lori Chronis, Betty Grimes, Carolyn C. Hoppe, and Stephen H. Waterman (Los Angeles County Department of Health Services, Los Angeles, CA); Brenda K. Boner, Rose A. Kelley, Lewis B. Lefkowitz, Jr, and Jo A. Taylor (Department of Health and Environment, State of Tennessee; Department of Preventive Medicine, Vanderbilt University School of Medicine, Nashville, TN); and Suzanne Gaventa and Arthur L. Reingold (CDC, Atlanta, GA).

criteria. Multivariate modeling was hypothesis-oriented and did not use stepwise variable selection methods.

Results

Surveillance and Identification of Case/Control Pairs

During 1986, 1,230 cases of invasive *H. influenzae* disease in children less than or equal to 60 months of age were reported from the five surveillance areas. ICD-9 discharge diagnosis code reviews were available for all surveillance areas except New Jersey. Of 408 cases reported by discharge diagnosis, 370 (90 percent) were detected by active surveillance. Day care status was determined for 1,081 (88 percent) of the 1,227 cases. A total of 358 cases attended day care facilities. Of these, 198 (57 percent) attended unlicensed day care facilities, and further information was not available for another 26 (7 percent), due to parental unwillingness to identify the facility. Parents or day care facility directors of 14 of the remaining 134 cases refused to participate. An additional four cases were deleted from the analysis because they occurred in facilities already enrolled as cases in the study, and we did not want to represent the same case facility twice in the analysis. These four facilities were not clearly different from other case facilities. Case and control pairs were obtained for 92 of the remaining 116 eligible cases (79 percent). The 24 eligible cases for whom controls were not available did not differ from the 92 included in the study by age, sex, race, or clinical syndrome, nor did their facilities differ by type or size or any of the variables associated significantly with case facilities in the univariate analysis. Sixty-nine case and control pairs were non-residential facilities (day care centers) and 23 pairs were residential facilities (day care homes). There were no significant differences in age, sex, race, or disease syndrome between day care attendees who were enrolled in the study and those who were not enrolled (Table 1).

Risk Factors

Univariate Analysis—Table 2 shows odds ratios of potential risk factors for *H. influenzae* disease in the analysis of all day care facilities regardless of age of case, enrollment, or type of facility. Separate analyses stratifying for age of case (less than or equal to 2 years of age, or older than 2 years of age), enrollment (less than or equal to 20 or greater than 20 children/day care facility) and type of day care facility (day care home vs day care center) did not identify other risk factors except as noted below.

TABLE 1—Characteristics of Day Care-Associated *H. influenzae* Cases

Characteristics	Cases Enrolled in Case-Control Study n = 92 (%)	Cases Not Enrolled* n = 266 (%)
Sex		
Male	49 (53)	130 (49)
Age (mo)		
0–6	13 (14)	50 (19)
7–11	25 (27)	71 (27)
12–23	29 (31)	83 (32)
24–35	10 (11)	37 (14)
36–59	15 (16)	25 (9)
Race		
White	63 (69)	181 (70)
Black	19 (21)	46 (18)
Asian/South Pacific	1 (1)	4 (2)
Hispanic	1 (1)	17 (7)
Unknown	8 (9)	18 (7)
Clinical syndrome (cultures positive)		
Cerebrospinal fluid (CSF) only	30 (33)	79 (30)
CSF + blood	29 (32)	94 (35)
Blood only	33 (35)	91 (34)

Note: Percent in parentheses.

*Because of unlicensed or unknown licensure status of facility, or refusal by parents or day care facility.

Age Range—The mean age range of children cared for in case-day care facilities was 67 months, while that in control centers was 83 months ($p = 0.023$). This difference was primarily a function of the difference in the mean age of the oldest child in the day care facility, since the mean age of the youngest child did not differ significantly between case and control day care facilities (due partly to the matching design).

Exclusion Policies—As shown in Table 3, control day care facilities were more likely to exclude children with diarrhea than were case day care facilities (OR 0.35, 95% CI: 0.14, 0.90). The two variables measuring exposure to feces were combined to form a third variable—liberal fecal policy (accepting non-toilet-trained children (Table 1) or not excluding children with diarrhea). The univariate odds ratio for liberal fecal policy was 3.3 (95% CI: 1.3, 8.3).

Business Status—Failure to use volunteers and for-profit status were highly associated variables with 15 percent of for-profit facilities using volunteers, compared with 39 percent of non-profit facilities ($p < 0.001$). Not using volunteers had an odds ratio of 2.6 and for-profit status had an odds ratio of 1.4.

Other Factors—Several other factors evaluated are listed in Table 2. Of note, use of a towel or handkerchief to wipe childrens' noses was associated with case day care facilities with an unadjusted odds ratio of 4.5.

Multivariate Analysis—A multivariate model was used to evaluate the significance of potential risk factors which were hypothesized to be related to disease and which, in the univariate analysis, showed either elevated (or reduced) odds ratios or associations ($p < .05$) with case facilities. Since two pairs of variables were closely related (age range/age of oldest child, and profit status/use of volunteers), only one of each pair was used in the multivariate analysis. The results of a multivariate model constructed with the variables of age range, profit status, liberal fecal policy and towel or handkerchief use are summarized in Table 4. Towel or handkerchief use is the most strongly associated variable (OR 5.5, 95% CI = 1.01, 30).

Discussion

During the course of this study, we found, as have others,[15] that many day care facility characteristics are closely associated. The matching technique and the association of many day care facility characteristics with the matching variables may have decreased our ability to identify risk factors in this study.

Use of towels or handkerchiefs to wipe childrens' noses was the most important risk factor for occurrence of *H. influenzae* disease in our multivariate model. This practice may facilitate transfer of the pathogen from child to child, since spread of *H. influenzae* organisms occurs through exchange of oropharyngeal secretions. The towel-handkerchief variable may represent a direct means of person-to-person spread of *H. influenzae*. *H. influenzae* organisms adhere to human nasopharyngeal mucous.[16] Murphy, *et al*, recently found that nasal secretions of *H. influenzae* carriers had much higher concentrations of *H. influenzae* than cough or oropharyngeal secretions, and that *H. influenzae* could

TABLE 2—Risk of *H. influenzae* disease for Selected Characteristics of Day Care Facilities, Univariate matched analysis (unadjusted)

Characteristics	n* (sets)	Cases	Controls	OR	95% CI
Mean age range (years)	91	5.1	6.4	0.9	0.82, 0.98
Mean age of oldest child in day care facility (years)	91	6.2	7.6	0.9	0.82, 0.98
Mean age of youngest child in day care facility (months)	91	7.8	8.9	0.98	0.94, 1.02
Mean hours open	89	58	58	1.0	0.87, 1.14
Mean number enrolled	92	70	64	1.0	1.00, 1.01
Mean number of staff	91	11	10	1.02	0.98, 1.06
Mean tuition ($/week, full time)	85	55	52	1.01	0.99, 1.02
Mean # children in class	82	9.7	9.8	1.0	0.93, 1.07
Mean # children in class 0–23 months old	65	5.5	5.3	1.04	0.91, 1.19
Mean age range in class	89	24	27	.99	0.97, 1.01
Mean age of youngest in class	89	13	13	.99	.94, 1.05
Mean age of oldest in class	89	37	40	.99	0.97, 1.01
% run for profit vs. nonprofit	92	63	57	1.37	0.72, 2.62
% using volunteers	88	16	34	0.38	0.19, 0.80
% accepting children for 1 day	92	33	48	0.54	0.29, 1.00
% accepting children not toilet trained	82	90	83	2.0	0.75, 5.33
% using disposable tissues	91	90	95	0.5	0.15, 1.66
% using towels or handkerchiefs**	83	10.9	2.4	4.5	0.97, 21

*analysis performed only on sets for which specific data were available.
**all facilities using towels or handkerchiefs also reported using disposable tissues.

TABLE 3—Exclusion Policies of Case and Control Day Care Facilities

Policy	n sets	case (%)	control (%)	odds ratio	95% CI
No policy	89	6.7	4.5	1.7	0.4, 7.0
Exclude fevers	92	86	90	0.64	0.25, 1.64
Exclude runny nose	90	10	10	1.0	0.29, 3.45
Exclude cough	89	12	19	0.54	0.21, 1.35
Exclude diarrhea	89	75	88	0.35	0.14, 0.90
Exclude vomiting	89	91	97	0.38	0.10, 1.4

TABLE 4—Multivariate Model (82 sets)

	OR	95% CI
Age range	.90	.80, 1.01
Profit (vs non-profit)	2.1	.93, 4.9
Towel/handkerchief use	5.5	1.01, 30
Liberal fecal policy	2.3	0.78, 6.9

survive for up to 48 hours on fomites such as water-absorbent towels.[17] Our study provides evidence that the route of transmission hypothesized by the laboratory studies of Murphy, *et al*, is important in day care facilities.

Several organizational characteristics were associated with case day care facilities in the univariate analysis. Case day care facilities were more likely to have a narrower age range and younger age of eldest child than control day care facilities. The age of youngest child in the day care facility did not differ. When day care centers were analyzed separately, these risk factors were still present. The same trends were present for day care homes, but were less stable. Although the association of narrower age range with risk of disease may be related to having more children of a younger age (and therefore at higher risk of *H. influenzae* disease) in a case day care facility, we could not evaluate this possibility since information on age of every child in each day care facility was not collected. However, neither number of classrooms nor number of children <23 months of age in each classroom differed between case and control day care facilities.

Slightly more control day care facilities than case day care facilities advised or required vaccination for *H. influenzae*, but the difference was minimal (42–44 percent vs 34–40 percent, depending on age of the child). It is unlikely that this difference accounted for lack of disease in control day care facilities because of the small size of the difference, and because more than two-thirds of all cases occurred in children too young to receive the vaccine (less than 18 months of age). The trend toward a more active vaccination policy, however, may represent better awareness of the disease and means of prevention.

Liberal fecal policy (evaluated because colonization with some enteric bacteria with cross-reacting antigens may be important in *H. influenzae* disease),[18] profit status, and age range were less strongly associated with case day care facilities. These variables were all related and may be surrogates for other more specific risk factors. For instance, profit status was also associated with hepatitis A incidence in the study of day care facilities in Arizona.[15] Although several other day care facility characteristics are likely to be associated with whether the facility is run as a for-profit or non-profit organization, it was not clear (from either this or the previous analysis of hepatitis A in day care facilities) how this could relate causally to incidence of *H. influenzae* disease.

Although a role for use of towels or handkerchiefs to wipe noses in transmission of infectious diseases has been suggested previously, this is the first study to identify it as a risk factor in the day care setting. Additional measures must be taken to eliminate this practice. Several other variables less strongly associated with risk of disease included liberal fecal policy, profit status, use of volunteers, and age range. They may be surrogates for more specific risk factors which remain to be identified.

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NIH Consensus Panel Issues Report on IVIG

Earlier this year, the National Institute of Allergy and Infectious Diseases and the Office of Medical Applications of Research of the National Institutes of Health convened a Consensus Development Conference on *Intravenous Immunoglobin: Prevention and Treatment of Disease*.

The consensus panel has now issued its report containing conclusions and recommendations from this conference.

Single copies of the report may be obtained from: William H. Hall, Director of Communications, Office of Medical Applications of Research, NIH, Bldg. 1, Rm 259, Bethesda, MD 20892; (301) 496-1142.

OMAR will also provide multiple copies of a two-page summary statement of the conference findings to organizations that wish to distribute copies to its members.

Injury and Disability in Matched Men's and Women's Intercollegiate Sports

Richard R. Lanese, PhD, Richard H. Strauss, MD, Daniel J. Leizman, MD, and Ann Marie Rotondi, MA

Abstract: Eight matched men's and women's intercollegiate varsity teams were studied prospectively for one academic year to determine the incidence of athletic injury and resulting disability. Sports in which both men and women participated in a comparable manner were chosen: basketball, fencing, gymnastics, swimming, tennis, indoor track, outdoor track, and volleyball. Men (232) and women (150) were injured at comparable rates, 42 percent versus 39 percent. When adjusted for exposure time, seven of the eight sports continued to show similar injury rates. Women gymnasts, however, experienced .82 injuries per 100 person-hours of exposure as compared to .21 injuries for the men ($p = .0001$). Disability was greater in women gymnasts, 7.44 days per 100 person-hours versus 1.15 days for men ($p = .0004$). Percent of season lost to injury was also greater for women gymnasts. Types and sites of injury were similar for men and women, with sprains and strains accounting for over half of all injuries. We found no evidence for gender differences in matched sports except for gymnastics, in which technically diverse events may have accounted for the differences observed. (*Am J Public Health* 1990; 80:1459–1462).

Introduction

Women's participation in intercollegiate athletics has increased dramatically in recent years. Several studies comparing injury rates in men and women participating in similar sports have reported conflicting results.[1–3] Some studies have noted higher injury rates in female gymnasts and basketball players.[4–7] However, few studies have been conducted on mature men and women.[8–10]

In view of the conflicting results of previous studies and the paucity of information on mature athletes, the present study was undertaken to provide injury data on men and women engaged in similar sports while controlling for the length of time each athlete was exposed to injury. Earlier studies have rarely controlled for exposure time. If gender differences are observed, a larger study may be designed to investigate possible risk factors such as technical differences within a sport or biological differences associated with injury.

Methods

The study was conducted as a prospective cohort study for the period of one academic year at a large university. Eight matched men's and women's varsity sports were studied. These were sports in which both men and women participated in a comparable manner and included basketball, fencing, gymnastics, swimming, tennis, indoor track, outdoor track, and volleyball.

One of the authors (A.M.R.) visited each team weekly to record the appropriate injury information for the week. Injury data came from the medical records kept by the team physicians and by certified and student athletic trainers.

An athletic trainer or student trainer was present at each team practice and competition and was responsible for collecting the injury and disability data. Injuries not attributable to the team sport, as well as all illnesses, were excluded from the study.

We defined an injury as a traumatic medical problem due to sports participation and resulting in loss of time from practice or competition. When an injury occurred during practice or competition and the athlete was unable to complete the day's activity, one-half day of disability was assigned. One whole day of disability was assigned for each full day of practice or competition lost. On weekends, teams usually had only one day off. For example, if an athlete was injured on Saturday, had no scheduled practice on Sunday, and returned to practice on Monday, half of Sunday was assigned as injury time off. Disability days were calculated for each athlete.

We studied each team from the first official day of practice until the last day of regular competition or practice. Hours of practice and competition were estimated by either direct observation or interviews with coaches and athletes, or a combination of both methods. This provided indices independent of differences in exposure time. Exposure time was calculated for each athlete and for each sport by subtracting disability days as well as other absences from the total time available for practice and competition.

To compare injuries between men and women, we utilized the following indices: 1) injuries per 100 persons; 2) injuries per 100 person-hours of exposure; 3) disability days per 100 person-hours of exposure; 4) percent of season lost to injury; and 5) percent of persons injured. These indices were computed for each athlete and each team. Because injuries were not normally distributed and, more than likely, not independent, the nonparametric Wilcoxon Rank Sum statistic,[11] corrected for ties, was employed to test gender differences within each sport and for all sports after assigning an injury rate to each athlete. This procedure was used for two indices: injuries per 100 person-hours of exposure, and percent of season lost to injury. The Mantel-Haenszel statistic[12] provided a test of difference in the proportion of men and women injured over all sports together. The relative risk of injury and its 95% confidence interval was estimated for males using the direct precision-based method.[13]

If the true proportions of male and female athletes who are injured are in the range of 40 percent to 60 percent, the present study sample has approximately an 80 percent chance of detecting a 15 percent difference when alpha is at .05.

Types of injury and anatomical sites were also determined. No practicable statistical methods of analyses were available to evaluate gender differences for these parameters.

Results

The study cohort was composed of 232 men and 150 women. Table 1 shows the frequency and percent injured for

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From the Department of Preventive Medicine, Ohio State University College of Medicine, Columbus. Address reprint requests to Richard R. Lanese, PhD, Professor of Preventive Medicine, Ohio State University, 320 West 10th Avenue, Columbus, OH 43210. Dr. Strauss is Associate Professor of Preventive Medicine and Internal Medicine, and team physician; Dr. Leizman was a medical student and Ms. Rotondi was a graduate student, all at Ohio State University, Columbus. This paper, submitted to the Journal September 5, 1989, was revised and accepted for publication April 17, 1990.

TABLE 1—Injuries among Male and Female Athletes in Eight Matched Intercollegiate Sports during One Year at a Large University

Sport	Gender	Persons on Team	Persons Injured (%)	Number of Injuries	Injuries per 100 Persons
Basketball	M	12	10 (83)	18	150
	F	14	11 (79)	19	136
Fencing	M	18	5 (28)	5	28
	F	6	3 (50)	3	50
Gymnastics	M	24	23 (96)	52	217
	F	11	11 (100)	30	273
Swimming	M	36	14 (39)	21	58
	F	21	8 (38)	8	38
Tennis	M	12	5 (42)	6	50
	F	11	3 (27)	4	36
Track, Indoor	M	57	21 (37)	28	49
	F	37	12 (32)	16	43
Track, Outdoor	M	56	9 (16)	11	20
	F	37	4 (11)	4	11
Volleyball	M	17	11 (65)	13	76
	F	13	6 (46)	7	54
All Sports	M	232	98 (42)*	154	66
	F	150	58 (39)	91	61

*Relative risk = 1.25 (CI 0.88, 1.43).

each of the eight matched sports. While the percentage of athletes injured varied considerably between sports, within sports there appear to be only small and inconsistent gender differences. For all sports combined, where 42 percent of the men and 39 percent of the women were injured, the Mantel-Haenszel χ^2 was 0.53 ($p \approx .50$). The relative risk, adjusted for between sport differences, was 1.25 with a 95% confidence interval of 0.88 to 1.43.

The rate of injury expressed as injuries per 100 persons was similar for all sports combined, 66 injuries per 100 persons for men, and 61 per 100 persons for women. Both men and women participating in gymnastics appear to have higher rates of injury than athletes in the other sports. All but one of the 35 gymnasts reported at least one disabling injury during the season. Women gymnasts were injured at a somewhat higher rate than men—273 versus 217 per 100 persons, respectively. When the incidence of injuries is expressed in terms of exposure time, women gymnasts sustained 0.82 injuries per 100 person-hours of exposure, while men incurred only 0.21 injuries (Table 2). This difference was significant by the Wilcoxon Rank Sum Test ($p = .0001$). There were no gender differences observed for the seven other sports nor for all sports combined.

Table 2 also summarizes disability days for men and women by sport. Women gymnasts had the highest rate of disability, 7.44 days per 100 person-hours of exposure, while men gymnasts had only 1.15 disability days, a sixfold difference ($p = .0004$). We were unable to detect gender differences in any of the remaining seven sports and, more importantly, when all eight sports were combined. For all sports, both men and women missed 1.4 days for every 100 hours of exposure, a period approximated by five to six weeks of practice and competition.

The types and sites of injury occurring during the period of study appear to be similar for both men and women. Strains and sprains were clearly the most common types of injury (Figure 1), accounting for over half of all injuries for both men and women.

The most frequent site of injury was foot and ankle (Figure 2). There appears to be little difference between sexes with respect to site. However, on closer inspection, injuries to the upper extremity (shoulder, upper arm, hand, wrist, elbow, and forearm) occurred more frequently in men (27 percent of injuries) than in women (13 percent of injuries).

Discussion

We relied on the medical records as kept by the team physicians and athletic trainers. The data were reviewed and accumulated once each week by one of the authors. We assumed that the reporting of injuries would be equally reliable across all sports and gender.

The present study involved 382 men and women over a period of one year at a single university. Thus, the question of repeatability of the findings from year to year, as well as the generalizability of the findings from one setting to another, may be questioned. While the total number of athletes studied was fairly large, the number of men and

TABLE 2—Injury and Disability Rates among Male and Female Athletes in Eight Matched Intercollegiate Sports during One Year at a Large University

Sport	Gender	Mean Hours of Exposure*	Injuries per 100 Person-Hours	(p)**	Disability Days	Disability Days per 100 Person-Hours	(p)**
Basketball	M	337	0.45	(.90)	65.5	1.62	(.48)
	F	283	0.48		69.0	1.74	
Fencing	M	278	0.10	(.27)	21.5	0.43	(.47)
	F	273	0.18		3.5	0.21	
Gymnastics	M	1020	0.21	(.0001)	281.5	1.15	(.0004)
	F	365	0.82		274.0	7.44	
Swimming	M	469	0.12	(.67)	241.5	1.43	(.77)
	F	488	0.08		63.5	0.62	
Tennis	M	305	0.16	(.37)	88.5	2.42	(.37)
	F	375	0.10		23.5	0.57	
Track, Indoor	M	266	0.18	(.44)	326.0	2.15	(.33)
	F	281	0.15		51.0	0.49	
Track, Outdoor	M	151	0.13	(.56)	143.5	1.70	(.52)
	F	92	0.12		71.5	2.11	
Volleyball	M	408	0.19	(.93)	52.0	0.75	(.48)
	F	353	0.15		22.5	0.49	
All Sports	M	397	0.18	p = .99	152.5	1.44	p = .49
	F	305	0.21		72.3	1.38	

*adjusted for missed practice and competition
**probability of difference between groups by Wilcoxon Rank Sum test

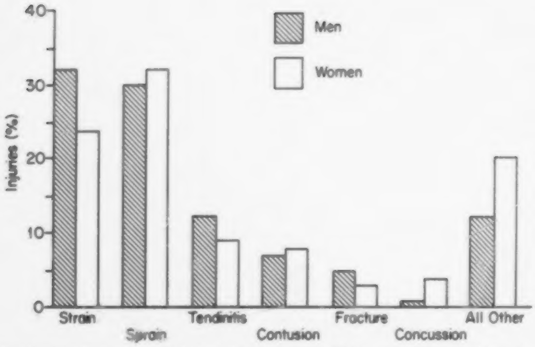

FIGURE 1—Percent of Injuries by Type among Male and Female Athletes in Eight Matched Intercollegiate Sports during One Year at a Large University

women in each paired sport was relatively small, resulting in low statistical power for detecting sport-specific gender differences. While gender differences in gymnastics were sufficiently large to overcome this problem, there may be differences in other sports that we were unable to detect.

Our estimates of exposure time for each athlete are crude, but should be useful when comparing gender differences for a single sport. For example, in basketball, men and women play the game in a similar manner and probably have comparable periods of activity and inactivity. To the extent that this is true, any gender difference in observed rate will be free of bias from exposure difference in a given sport.

We expected modest differences in exposure times between women and men in the same sport. However, we did not anticipate the size of the differences observed. For the 16 teams, exposure times ranged from 92 hours for women's outdoor track to 1,020 hours for men's gymnastics. It is clear from Table 2 that within-sport and between-sport comparisons of injury must take these exposure differences into account. Gymnastics reflects the greatest exposure disparity between women and men, 365 hours and 1,020 hours, respectively. Yet, women report a higher incidence of injuries per 100 persons.

We found little difference between men and women in terms of injury or disability in the eight sports except for gymnastics. Other investigators[2,10,14] as well, have reported differences in injury rates and types of injury in male and female gymnasts, with females showing somewhat higher rates. The differences that we observed are largely a result of expressing injuries in terms of exposure time. Only small differences between the sexes were observed when injuries were expressed per 100 persons or as percent of persons injured (Table 1).

When we expressed disability as the percent of season lost for each athlete, the median for women gymnasts, 8.21 percent, was significantly higher than the median for men, 3.15 percent ($p < .01$). No other team lost as high a percentage of their season due to disability. Ten of the 14 teams had medians at zero percent.

It is generally accepted that men gymnasts use predominantly upper body musculature and women use predominantly lower body musculature. This is consistent with the upper extremity injury difference noted earlier. The injury and disability differences that have been reported in this paper and elsewhere may be more a consequence of the different types of events and apparatus than of gender. Women gymnasts participate in four events and men in six. Only two of the events utilize the same apparatus: floor exercise and vault.

It is interesting to note that, while women and men gymnasts comprised only 9 percent of the athletes followed during the year of study, they accounted for almost 34 percent of all injuries. The apparent excess of injuries in gymnasts may be a result of the special requirements of the sport. Injuries that might be regarded as inconsequential and not reported by athletes in other sports such as football may impair the performance of the gymnast.

The main finding of this paper suggests that, when men and women compete in sports that are technically similar, the

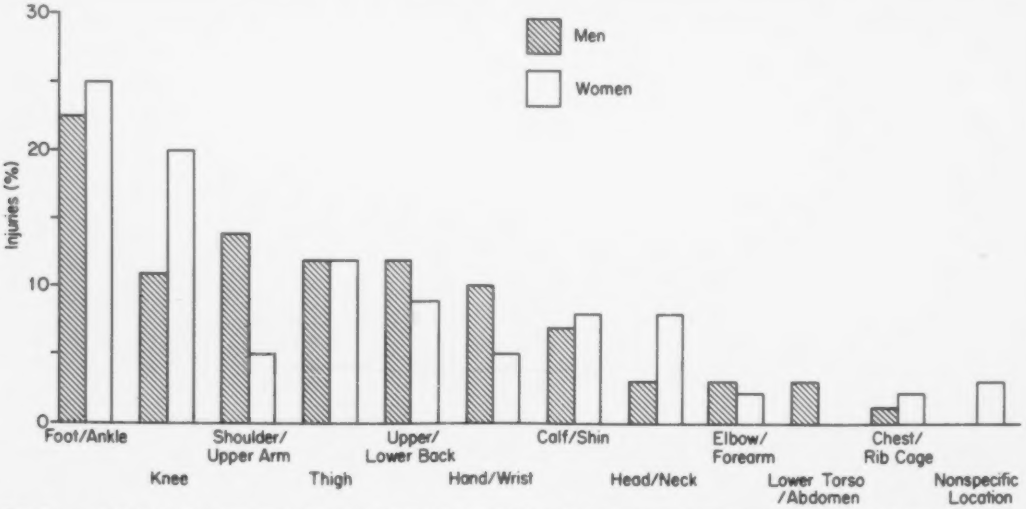

FIGURE 2—Percent of Injuries by Site among Male and Female Athletes in Eight Matched Intercollegiate Sports during One Year at a Large University

rates of injury are approximately equal. When they compete in sports that are technically different, there is little reason to expect injury rates to be the same. This may account for our single positive finding—that men and women gymnasts differed in injury rates.

ACKNOWLEDGMENT

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National Conference on Cancer Nursing: Call For Abstracts

Nursing professionals are invited to submit abstracts for poster presentations for the American Cancer Society's Sixth National Conference on Cancer Nursing. The conference, "From Primary to Tertiary Care: Cancer Across the Life Span," will be held July 25–27 in Seattle, Washington.

Within the theme, abstract topics may be related to clinical practice, education, research, and nursing administration. The deadline for abstract submission is February 1. For additional information, contact: Terri Ades, RN, Coordinator for Nursing Programs, ACS National Office, 1599 Clifton Rd, NE, Atlanta, GA 30329; (404) 329-7655.

This conference is the sixth in a series of national nursing conferences which are held every four years, and whose purpose is to improve the quality of cancer care provided to individuals and their families. The conference program is directed at nurse generalists, primary care nurses, oncology nurses, and other members of the health profession whose practice is devoted to providing care to individuals and their families throughout the life span. The conference is sponsored by the American Cancer Society, in cooperation with the American Nurses' Association Council on Primary Health Care Nurse Practitioners, Association of Pediatric Oncology Nurses, and the Oncology Nursing Society.

Coding External Causes of Injury (E-Codes) in Maryland Hospital Discharges 1979–88: A Statewide Study to Explore the Uncoded Population

BARUCH MARGANITT, MD, MPH, ELLEN J. MACKENZIE, PHD, GORDON S. SMITH, MBCHB, MPH,
AND ANNE M. DAMIANO, SCM

Abstract: We examined the trends in hospital discharge E-coding in Maryland over a 10-year period. The overall proportion of E-coded discharges has increased from 40 percent in 1979 to 55 percent in 1988. E-coding was lower in the severely injured, the elderly, and patients with long hospital stays. Our findings demonstrate that E-code reporting varies because of the limited number of data fields available for coding of discharge diagnoses. Universal, systematic reporting of E-codes in hospital discharge data is essential if these data are to provide critically needed information about nonfatal injuries. Hospital discharge data formats should contain separate fields for E-codes and the use of these codes, we believe, should be mandated. (*Am J Public Health* 1990; 80:1463–1466).

Introduction

Efforts to examine trends in the incidence of injury, patterns of acute care, and patient outcomes have generally required substantial investments in case ascertainment and data collection. The increasing availability of computerized statewide hospital discharge data offer new opportunities for timely and efficient data collection and policy analysis of issues related to the prevention of injuries and trauma care delivery.[1–8] The recent development of a computerized conversion table that translates ICD-9CM coded discharge diagnoses into Abbreviated Injury Scale (AIS) scores has led to new applications of these databases, providing population-based data specific to the severity of injuries.[8–11] Currently, 28 states require that uniform hospital discharge data be collected on all acute care hospitalizations.[8]

The lack of consistent coding of the external cause of injury limits the usefulness of these databases.[12,13] Although a classification of external causes exists within the structure of the International Classification of Diseases (ICD E-codes-),[17] hospitals vary in their use of these codes. Accurate data on the causes of nonfatal injuries and a means to measure their trends are not available in the United States. Underreporting of E-codes in hospital discharge abstracts has been recognized as a major limitation in their use for these purposes.[12–15]

To better understand factors responsible for incomplete E-coding, we analyzed 10 years of Maryland hospital discharge data. The specific objectives of the study were: to examine time trends in the level of E-coding in the statewide population-based hospital discharge data; and to determine demographic and clinical characteristics of the uncoded population (the missing E-code population), in order to identify strategies for improving reporting of E-codes.

Methods

Uniform hospital discharge abstract data for the state of Maryland during the 10 years from 1979 to 1988 were used to examine trends in E-coding over time. All 52 non-federal acute care hospitals in Maryland are mandated by law to submit uniform data to the State's Health Services Cost Review Commission (HSCRC). External causes of injury are reported as ICD-9CM E-Codes and can be included in the database as one or more of five possible discharge diagnoses.

The Study Population consisted of the 349,487 hospital discharges in the state of Maryland with a principal diagnosis of traumatic injury. Traumatic injuries were defined as those injuries coded with ICM-9CM diagnoses codes 800–959 excluding: late effects (905–909), foreign bodies (930–939), and early complications (958). Poisonings (960–979), toxic effects (980–989), other and unspecified effects of external causes (990–995), and complications of medical and surgical care (996–999) were also excluded. E-coding status of each discharge abstract was defined as one of the following three categories based on existence of an E-code and availability of ICD-9CM code recording space: 1) E-coded discharge, 2) Uncoded discharge with no coding space available (no E-code, not codable), 3) Uncoded discharge with coding space available (no E-code, codable). Using the 1988 data, coding status was examined by age of the patient, injury severity, body region of the injury assigned as the principal diagnosis (head-neck/extremities), length of hospital stay, and level of hospital care, i.e.: 1) Specialty Referral Trauma Centers (SRTC) (One adult and one pediatric); 2) Regional Trauma Center (RTC) (nine centers); 3) Community Hospital (Com Hosp) (41 hospitals). Injury Severity was assessed by the Injury Severity Score (ISS)[16] and was assigned to each discharge using the ICDMAP software[9,10] that maps ICD-9CM coded diagnoses into Abbreviated Injury Scale (AIS) scores and ISS. E-code status was examined for the 10 years 1979–1988.

Results

The proportion of E-coded trauma discharges increased over the study period by 38 percent, from 40 percent in 1979 to 55 percent in 1988 (Figure 1). Over the same period, however, the proportion of uncoded discharges that had vacant space available for additional codes (no E-code, codable) decreased from 98 percent in 1979 to 53 percent in 1988.

Increasing age was associated with a decrease in the percent of E-coded discharges as well as an increase in the proportion of not codable discharges (Figure 2). In 1988, the year with the highest coding rate, 73 percent of the 0–5 age

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From the Health Services Research and Development Center and Injury Prevention Center, Johns Hopkins School of Hygiene and Public Health. Address reprint requests to Baruch Marganitt, MD, MPH, Health Services Research and Development Center, Johns Hopkins School of Hygiene and Public Health, 624 N. Broadway Street, Baltimore, MD 21205. This paper, submitted to the Journal January 16, 1990, was revised and accepted for publication April 20, 1990.

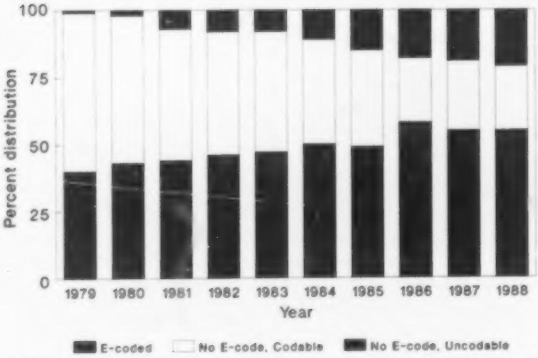

FIGURE 1—E-Coding Status by Year, Maryland Trauma Discharges, 1979–88

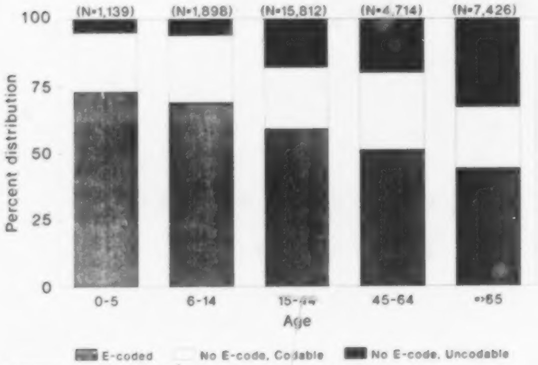

FIGURE 2—E-Coding Status by Age, Maryland Trauma Discharges, 1988

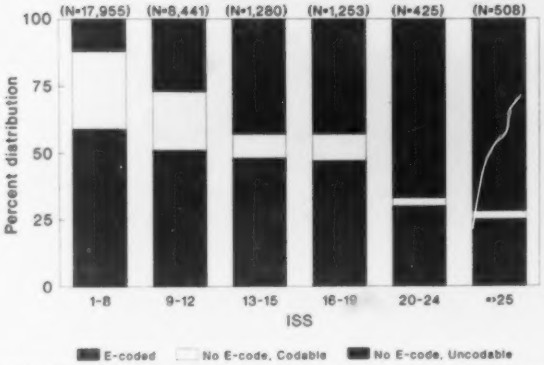

FIGURE 3—E-Coding Status by Injury Severity Score, Maryland Trauma Discharges, 1988

group discharges were E-coded as compared to 44 percent in the ≥65 age group. However, in age group 0–5, 18 percent of the uncoded discharges were not codable, as compared to 59 percent in age group ≥65. This differs from earlier study years; in 1979, for example, the proportion of E-coded discharges ranged from 44 percent in 0–5 age group to 40 percent in ≥65, while the percent of uncoded discharges that were not codable ranged from 0–1 percent.

As the Injury Severity Score (ISS) increased the proportion of E-coded discharges decreased and the proportion of uncoded discharges that were not codable increased (Figure 3). In 1988, the proportion of E-coded discharges ranged from 59 percent of ISS group 1–8 to 25 percent of ISS ≥25. However, in ISS group 1–8, 29 percent of the uncoded discharges were not codable, as compared to 92 percent of the uncoded discharges of ISS group ≥25. In the early years of the study period, E-coding status did not vary with ISS; for example, in 1979, the proportion of E-coded discharges ranged from 42 percent of ISS 1–8 to 39 percent of ISS ≥25; out of the uncoded discharges the percent of not codable discharges, however, was very low ranging from 0–2 percent.

Table 1 compares the 1988 E-coding status within ISS subgroups between the two most prevalent body-region categories: head/neck and extremities. In both groups the proportion of E-coded discharges decreased with increasing ISS. However, the proportion of coded discharges in the extremities group was lower than the head-neck group in all ISS categories, with more notable differences in the most severe groups (ISS 16 and higher). There were also marked differences between the head/neck and extremities groups in the proportion of uncoded discharges that were not codable. For example, in the ISS 16–19 Head/Neck group 47 percent of the discharges were uncoded, and 72 percent of them were not codable. This is compared to discharges with a principal injury to the extremities of which 76 percent were uncoded, and of them 95 percent were not codable.

In 1988, 64 percent of the 1 day LOS group were E-coded, as compared to 25 percent of the 29–56 days, and 12 percent of LOS ≥57 days (Figure 4). Of the uncoded discharges the proportion that were not codable increased from 17 percent of the uncoded discharges in the 0–1 day LOS group to 96 percent of the uncoded discharges in the ≥57 days LOS group.

E-coding increased over the study period in all levels of care, however trauma centers coded a higher proportion of discharges. For example, in 1988 over 60 percent of trauma center discharges were E-coded as compared to 47 percent of the discharges from community hospitals (Figure 5). In community hospitals, 70 percent of the uncoded discharges were codable, as compared to 28 percent and 7 percent of the uncoded discharges in Regional Trauma Centers and Specialty Referral Trauma Centers, respectively.

Discussion

This study demonstrates that the current structure of many statewide hospital discharge data systems restricts the use of E-codes for specific subgroups of patients. In Maryland, reporting of E-codes is not required and is done inconsistently across hospitals and patients. In many hospital discharge data systems, including HCFA's UB-82, E-codes can only be included in the discharge abstract as one of five discharge diagnoses.[13] Since ICD E-codes are not used as a basis for reimbursement, they receive low priority in the selection process of the five ICD-9CM codes reported.[14] Thus E-coding rates are lowest among selected groups of patients that are typically characterized by large number of discharge diagnoses, i.e. severely injured patients having multiple injury-related diagnoses, the elderly patients who are likely to have chronic diseases existing prior to the trauma and are more prone to complications leading to additional discharge diagnoses, and the patients with long hospital stays either due to severe injuries or complications. Injuries, preexisting chronic conditions and complications all take precedence

TABLE 1—E-Coding Status by Body-Region of Principal Diagnosis and by ISS, Maryland Trauma Discharges, 1988

	%E-Coded		% No E-Code Codable		% No E-Code Uncodable		Total	
ISS	Head/ Neck	Extremities	Head/ Neck	Extremities	Head/ Neck	Extremities	Head/ Neck	Extremities
1–8	65	54	20	38	15	8	100	100
9–12	61	48	12	25	27	27	100	100
13–15	52	44	3	8	45	48	100	100
16–19	53	24	13	4	34	72	100	100
20–24	26	22	2	5	72	73	100	100
≥25	31	12	3	2	66	86	100	100

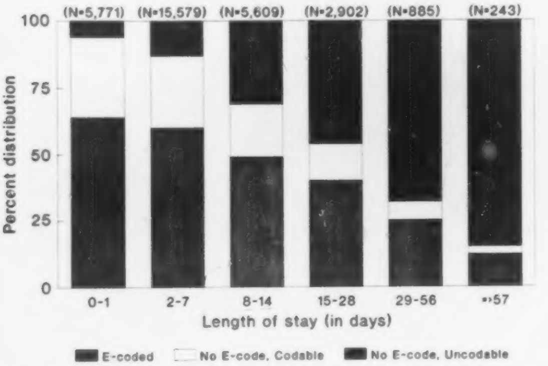

FIGURE 4—E-Coding Status by Length of Hospital Stay, Maryland Trauma Discharges, 1988

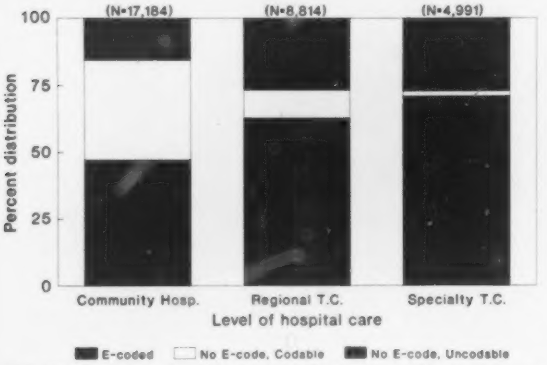

FIGURE 5—E-Coding Status by Level of Hospital Care, Maryland Trauma Discharges, 1988

over cause in coding since they may influence the rate at which hospitals are reimbursed.[14]

In Maryland, the proportion of E-coded trauma discharges has increased from 40 percent in 1979 to 55 percent in 1988. This is likely due to increasing awareness and interest in injury control and Emergency Medical Services (EMS) in the state. Of the uncoded discharges, however, the proportion that were codable has decreased from 98 percent in 1979 to 53 percent in 1988. This trend was most marked in the selected groups of patients described above. This finding suggests that while there is an increasing awareness of the importance of E-coding in the state, there is less potential for including an E-code in the hospital discharge data. This is likely due to the increased attention paid to characterizing the full extent and nature of the injury, complications, and comorbidities and utilization of all five ICD-9CM coding spaces.

Simply increasing the number of fields for recording diagnoses, however, is not an adequate solution. In California, for instance, where a maximum of 25 hospital discharge diagnoses can be recorded,[18] only 40 percent of the trauma discharges in 1983 have been E-coded (MacKenzie EJ, unpublished data). Rather, it will be necessary to mandate E-coding and provide data fields separate from those available for recording diagnoses, preexisting medical conditions, and complications. Mandatory E-coding legislation or regulation, as initiated in New York, California, Wisconsin, and Washington State,[13] and more recently in Rhode-Island and Vermont, provide the means to ensure better compliance with E-coding practice. As has been demonstrated in this study, however, it is essential that separate fields for recording of E-codes be created otherwise it may not be recorded on computerized data bases for those injury cases with multiple diagnoses. There is a need for two E-coding fields in order to adequately characterize the external cause and intent of the injury.[19] Monitoring trends in E-coding rates in these states will be important to assess the extent of compliance and accuracy. Mandating that trauma discharges be E-coded will only be effective if hospital coders are properly trained.[13] Computer software recently developed by the Centers for Disease Control should assist in simplifying E-coding and increase in reliability.[20] In addition, training manuals and courses for E-coding are now available.[19]

Until such time as E-codes are uniformly recorded for all trauma discharges, caution must be exercised when extrapolating results by cause of injury based on only those discharges that are E-coded. This study shows systematic underreporting of external causes of injury in specific trauma patients groups such as the elderly, the severely injured, etc. At the very least, extrapolations should only be made within strata of the population defined by age of the patient and the nature and severity of the injuries sustained. It is also important to recognize that patterns of E-coding may vary over time, by geographic region and by hospital type and coding practices. Gathering cause of injury data from different databases requires attention to the scope and direction of E-code reporting bias that each system may introduce.

Hospital discharge data represent an important and usually the only source of population-based data on serious, nonfatal injuries. Systematic, uniform, and reliable reporting of E-codes in these databases would provide essential information for monitoring trends and patterns of injury, defining subgroups of the population at high risk of injury, and in developing and evaluating injury prevention programs.

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Benzodiazepine Tranquilizers and the Risk of Accidental Injury

Gerry Oster, PhD, Daniel M. Huse, MA, Shelley F. Adams, MPA,
Joseph Imbimbo, BA, and Mason W. Russell, MAPE

Abstract: To determine whether benzodiazepine tranquilizers increase the risk of accidental injury requiring medical attention, we used pharmacy claims submitted to a large third-party payer to identify 4,554 persons who had been prescribed these agents and a matched control group of 13,662 persons who had been prescribed drugs other than benzodiazepines. We then used diagnoses recorded on claims submitted by medical care providers to identify all accident-related care received by these persons during three months before their first-observed prescription for a benzodiazepine or nonbenzodiazepine agent, respectively, and six months subsequently. We found accident-related care was more likely among persons who had been prescribed benzodiazepines; among these persons, the probability of an accident-related medical encounter was higher during months in which a prescription for a benzodiazepine had recently been filled compared to other months; and persons who had filled three or more prescriptions for these agents in the six months following initiation of therapy had a significantly higher risk of an accident-related medical event than those who had filled only one such prescription. Approximately two-fold risks of accident-related care were found, after controlling for age, sex, and prior utilization. (*Am J Public Health* 1990; 80:1467–1470.)

Introduction

Benzodiazepine tranquilizers were introduced in the United States in 1962 and rapidly supplanted barbiturates as anxiolytic agents of first choice. While they are among the most widely prescribed classes of drugs, there is mounting evidence that their use impairs cognitive performance and psychomotor skills. Persons given benzodiazepines have been found to perform poorly in driving simulator tests relative to those receiving placebo.[1,2] In addition, several forensic studies[3–6] have found a greater than expected number of benzodiazepine users among persons fatally injured in motor vehicle accidents.

Three epidemiologic investigations of impairment related to the use of minor tranquilizers have been reported to date. A study conducted in England over a two-year period found that patients receiving benzodiazepines and other minor tranquilizers were nearly five times more likely to experience a serious motor vehicle accident than those who had not used these drugs.[7] Another study examined persons hospitalized for injuries resulting from a motor vehicle accident and found that persons identified as drivers who were at fault in their accidents were somewhat more likely to have filled a prescription for hypnotics or tranquilizers during the three-month period prior to the accident than those identified as passengers.[8]

The third epidemiologic investigation, a retrospective cohort study using claims data from a large health insurer,[9] compared accident-related medical care among persons who were prescribed benzodiazepines with that of a matched control group who had not been prescribed these agents. Persons prescribed benzodiazepines were more likely to have accident-related care, but also were reported to experience a greater number of nonaccident-related medical encounters, suggesting differences in care-seeking behavior rather than risk of accidental injury. Without data on medical utilization prior to therapy, the effect of differences in care-seeking behavior could not be determined.

To circumvent this difficulty, in the present study we employ a pretest-posttest design[10] to examine three key questions: 1) Are persons who are prescribed benzodiazepine tranquilizers more likely to require accident-related medical care? 2) Is there evidence of a temporal relation between prescriptions for these agents and the likelihood of an accident-related medical encounter? 3) Is there evidence of a relation between the number of such prescriptions and the risk of these encounters?

Methods

Data Source

Data for this study were obtained from the health care claims processing system of Master Health Plus™, a health insurance plan operated by Blue Cross and Blue Shield of Massachusetts with a membership in excess of one million persons. Plan members receive first-dollar coverage for physician services, inpatient and outpatient care, and prescription drugs, with a $5 copayment for provider services, a $4 copayment for prescription drugs ($3 for generics), and a $500 annual limit on reimbursement for mental health services. All claims for care are provider-submitted, and nondrug claims contain detailed diagnostic information in the form of physician-recorded ICD-9-CM codes. A complete record of medical utilization may be compiled for each plan member during his or her period of enrollment.

To be eligible for inclusion in the present study, members had to be enrolled continuously in the plan during the period April 1986–September 1987. Persons over 65 years of age or otherwise eligible for Medicare benefits were excluded from the study because Medicare claims were not available.

For our sample of persons who were prescribed benzodiazepines (benzodiazepine "users"), we selected all eligible plan members who had at least one pharmacy claim for a benzodiazepine tranquilizer (alprazolam, chlordiazepoxide, clorazepate, diazepam, lorazepam, or oxazepam, either singly or in combination with any other agent) during July 1986–March 1987, and had no claims for any of these agents between April 1986–June 1986. We considered the date of the first-observed pharmacy claim for a benzodiazepine agent to be the date of initiation of benzodiazepine therapy. A total of 4,554 benzodiazepine users were identified and included in the study.

From Policy Analysis Inc. (Oster, Huse, Russell); New England Institute for Healthcare Services Research (Adams); and Electronic Data Systems, Inc. (Imbimbo). Address reprint requests to Gerry Oster, PhD, Policy Analysis Inc., 1577 Beacon Street, Brookline, MA 02146. This paper, submitted to the Journal July 11, 1989, was revised and accepted for publication April 17, 1990.

Three persons who were not prescribed benzodiazepines (benzodiazepine "nonusers") were matched randomly to each user. Nonusers were selected from among those eligible patients who had no pharmacy claims for a benzodiazepine tranquilizer between April 1986–March 1987, were not related by blood or marriage to any patient who had a claim for such an agent during April 1986–March 1987, and had at least one claim for a nonbenzodiazepine agent during July 1986–March 1987. We considered the date of the first-observed nonbenzodiazepine drug claim to be the date of therapy initiation for each nonuser, and we matched nonusers to users on the basis of age, sex, and the calendar month during which therapy was initiated.

For each person in the sample, we then compiled all health care claims for three months prior to the date of initiation of therapy, and for six months subsequent to that date. Every nondrug claim was classified as either accident- or nonaccident-related on the basis of the physician-recorded ICD-9-CM diagnosis codes. Claims with any ICD-9-CM diagnosis code between 800-959 or 990-995 were considered to be accident-related; all other claims were considered nonaccident-related. Emergency visits were identified on the basis of accompanying information noting the type of service rendered (i.e., "emergency medical care").

Univariate Analysis

To examine the relation between benzodiazepine use and overall risk of accidental injury, we considered three outcomes: 1) the occurrence of any accident-related *emergency outpatient visit;* 2) the occurrence of any accident-related *hospital admission;* and 3) the occurrence of an accident-related *medical encounter of any type*. The frequencies of these events among benzodiazepine users and nonusers during the six-month period following initiation of therapy were determined.

To examine evidence of a temporal relationship between benzodiazepine use and risk of accidental injury, among users we compared the frequencies of accident-related medical encounters during months of presumed drug exposure to those during months of presumed nonexposure. A user was presumed to have been "exposed" to benzodiazepines during a given month if a pharmacy claim for one of these agents was observed in that month or the previous month.

Finally, to ascertain whether there is evidence of a relation between the number of benzodiazepine prescriptions filled and the risk of accidental injury, we compared rates of accident-related medical events among users according to their numbers of pharmacy claims for benzodiazepine agents during the six months following initiation of therapy. Two groups of users were defined—those who had only one pharmacy claim for a benzodiazepine agent and those who had three or more such claims—and we compared rates of accident-related medical events between them.

Comparisons of rates were expressed as ratios, or relative risks, and 95 percent confidence intervals were calculated by the Taylor-series method.[11]

Multivariate Analysis

To control for factors other than benzodiazepine use, we fitted a Cox proportional hazards model[12] and included as covariates age, sex, and the logarithms of total accident- and nonaccident-related medical care charges during the three months prior to therapy.

We used a similar proportional hazards model to compare the risk of accident-related medical encounters during months in which a user was presumed to have been exposed to benzodiazepines to all other months. Included in this model were a set of additional binary variables indicating, for each month during the six-month period of follow-up, whether a user was presumed to have been exposed to benzodiazepines in that month.

Finally, we used a similar model to compare outcomes between benzodiazepine users who filled three or more prescriptions for these agents and those who filled only one such prescription. Risks of each accident-related event were estimated as a function of the number of benzodiazepine prescriptions filled, age, sex, and the logarithms of total accident- and nonaccident-related medical-care charges during the three months prior to therapy.

Results

Profile of Benzodiazepine Users and Patterns of Benzodiazepine Use

Benzodiazepine users were more likely to be women than men (65.4 percent versus 34.6 percent), and more than one-half (53.5 percent) were less than 45 years of age. Approximately one-half (48.3 percent) of benzodiazepine users filled only a single prescription for one of these agents, 21.3 percent filled two prescriptions, 11.2 percent filled three prescriptions, and 19.2 percent filled four or more prescriptions over the six-month period of follow-up.

During the three-month period prior to the initiation of therapy, benzodiazepine users were twice as likely as nonusers (95 percent CI: .96, 4.44) to have at least one accident-related emergency outpatient visit (2.4 versus 1.2 per 1,000); 1.7 times as likely (95 percent CI: 1.11, 2.61) to have an accident-related hospital admission (7.2 versus 4.2 per 1,000); and 1.5 times as likely (95 percent CI: 1.29, 1.66) to have an accident-related medical encounter of any type (73.6 versus 50.2 per 1,000.

Accident-Related Care among Benzodiazepine Users and Nonusers

During the six-month period following the initiation of therapy, benzodiazepine users were significantly more likely than nonusers to require accident-related care (Table 1). Compared to nonusers, users were more than twice as likely to have had an accident-related emergency outpatient visit, about 1.6 times more likely to have experienced at least one accident-related hospital admission, and approximately 1.2 times more likely to have had at least one accident-related medical encounter of any type.

We examined rates of accident-related encounters among persons who had no accident-related care during the three-month period prior to the initiation of therapy. Unadjusted relative risks of accident-related events among benzodiazepine users were similar to those reported in Table 1.

TABLE 1—Risks of Accident-related Medical Encounters among Persons Who Were Prescribed Benzodiazepine Tranquilizers ("users") Compared to Those Who Were Not ("nonusers")

Accident-Related Medical Encounter	Monthly Rate per 1,000		Unadjusted Relative Risk (95% CI)	Adjusted Relative Risk* (95% CI)
	Users	Nonusers		
Any emergency outpatient care	1.06	.49	2.16 (1.35,3.50)	2.09 (1.27,3.42)
Any hospital admission	1.39	.88	1.58 (1.07,2.34)	1.27 (.84,1.90)
Any medical encounter	23.31	18.84	1.24 (1.14,1.35)	1.15 (1.05,1.26)

*Hazard ratio adjusted for age, sex, and pretreatment medical care utilization.

For emergency outpatient visits, the unadjusted relative risk was 2.18 (95 percent CI: 1.27, 3.77); for hospital admissions, it was 1.77 (95 percent CI: 1.13, 2.78); and for medical encounters of any type, it was 1.21 (95 percent CI: 1.10, 1.33). Controlling for age, sex, and health care utilization in the three months prior to the initiation of therapy did not change the excess risk of accident-related medical encounters among benzodiazepine users relative to nonusers appreciably (Table 1).

Accident-Related Care by Temporal Pattern of Benzodiazepine Use

Outcomes among benzodiazepine users were examined during 11,389 months of presumed drug exposure (i.e., months in which prescriptions for benzodiazepines had recently been filled) and 15,935 months in which they were presumed unexposed. A greater likelihood of accident-related medical encounters was observed during months of exposure (Table 2). Controlling for age, sex, and health care utilization in the three months prior to treatment, adjusted relative risks were somewhat lower than unadjusted ratios but remained excessive.

Accident-Related Care by Amount of Benzodiazepine Use

Benzodiazepine users who filled at least three prescriptions for these agents (n = 1,384) were more likely to have required accident-related care than those who filled only one such prescription (n = 2,200) (Table 3). Compared to users who filled one prescription, those who filled three or more prescriptions were about 2.4 times more likely to have had an accident-related emergency outpatient visit, and roughly 1.3 times more likely to have had an accident-related medical encounter of any type. The relative risk of an accident-related hospital admission was 1.5. Controlling for age, sex, and health care utilization during the three-month period prior to treatment did not change these risks substantially (Table 3).

Discussion

The results of this study suggest that the use of benzodiazepine tranquilizers increases the risk of accidental injury requiring medical attention. Emergency outpatient care, hospital admissions, and medical encounters of all types were more frequent among persons who were prescribed benzodiazepines than those who were not; among persons who were prescribed benzodiazepines, the likelihood of such care was higher during months in which such a prescription had recently been filled than during other months; and persons who filled three or more prescriptions for benzodiazepines over a six-month period had greater numbers of accident-related medical encounters than those who had filled only one such prescription. Increased risks of accident-related care remain even after controlling for the potentially confounding effects of age, sex, and prior differences in utilization.

TABLE 2—Risks of Accident-Related Medical Encounters among Persons Who Were Prescribed Benzodiazepine Tranquilizers during Months in Which Such a Prescription Was Recently Filled ("exposed") Compared to other Months ("unexposed")

Accident-Related Medical Encounter	Monthly Rate per 1,000		Unadjusted Relative Risk (95% CI)	Adjusted Relative Risk* (95% CI)
	Exposed	Unexposed		
Any emergency outpatient care	1.23	.94	1.31 (.63,2.70)	1.13 (.46,2.77)
Any hospital admission	1.84	1.07	1.73 (.91,3.27)	1.42 (.66,3.04)
Any medical encounter	31.87	17.19	1.85 (1.59,2.16)	1.28 (1.04,1.56)

*Hazard ratio adjusted for age, sex, and pretreatment medical care utilization.

TABLE 3—Risks of Accident-related Medical Encounters among Persons Who Filled Three or more Prescriptions for Benzodiazepine Tranquilizers ("scripts") Compared to Those Who Filled one such Prescription

Accident-Related Medical Encounter	Monthly Rate per 1,000		Unadjusted Relative Risk (95% CI)	Adjusted Relative Risk* (95% CI)
	Three or More Scripts	One Script		
Any emergency outpatient care	1.81	.76	2.38 (1.16,6.04)	2.64 (1.15,6.04)
Any hospital admission	1.57	1.06	1.48 (.70,3.13)	1.42 (.67,3.03)
Any medical encounter	27.70	21.52	1.29 (1.10,1.51)	1.30 (1.09,1.55)

*Hazard ratio adjusted for age, sex, and pretreatment medical care utilization.

We note that approximately one-third of the benzodiazepine users in our sample filled three or more prescriptions for these agents over a six-month period, and that these persons were at highest risk. Compared to persons who were not prescribed benzodiazepines, persons who filled three or more such prescriptions were 3.7 times (95 percent CI: 2.05, 6.68) more likely to have had an accident-related emergency outpatient visit (10.8 versus 2.9 per 1,000), 1.8 times (95 percent CI: 0.99, 3.21) more likely to have had an accident-related hospital admission (9.4 versus 5.3 per 1,000), and 1.5 times (95 percent CI: 1.30, 1.67) more likely to have had at least one accident-related medical encounter of any type (166.2 versus 113.0 per 1,000).

Our control group consisted of patients who received prescription drugs other than benzodiazepines, some of which may also affect the central nervous system. Our findings concerning the increase in risk of accidental injury associated with the use of benzodiazepines may therefore be conservative. Since we do not know whether persons who filled prescriptions for benzodiazepines in fact took any medication and, if so, how much they took and for how long, we may have overestimated drug exposure in our user group. Similarly, there may have been some exposure to these agents among our nonusers. While we excluded all persons with pharmacy claims for benzodiazepines from our nonuser group as well as those related by blood or marriage to these persons, we could not exclude those who had received these agents directly from their physician or those who had obtained these drugs illicitly. Our estimates of relative risk may therefore be conservative.

We considered the possibility that benzodiazepine users simply may be more accident prone than nonusers, due to differences in behavior, lifestyle, occupation, and so forth. We feel that this is unlikely, however; such factors could not account for the *temporal* association between benzodiazepine use and accident-related care that we observed among patients receiving these agents.

Since benzodiazepine tranquilizers often are prescribed for acute anxiety, we also considered the possibility that anxiety itself rather than the prescribed medication may cause users of these agents to be more accident-prone. Accordingly, we identified all persons who had claims for mental health services at any time during the 30 days prior to or following the initiation of benzodiazepine or nonbenzodiazepine therapy, but no such claims previously. When included as a covariate in our proportional hazards model, however, this variable was found to be uncorrelated with patients' risk of accident-related care, and was subsequently dropped. We note also that the results of controlled experiments lend credence to our finding that benzodiazepine use per se increases the risk of accident-related care.

We also considered the possibility that benzodiazepine users simply may be more likely than nonusers to seek medical treatment for minor injuries. Benzodiazepine users have been found to have higher rates of medical encounters unrelated to accidents than nonusers,[8] which suggests that users might be prone to seek professional care more often than nonusers. Again, the temporal association between the filling of prescriptions for benzodiazepines and accident-related encounters cannot readily be explained by differences in care-seeking behavior. Also, the excess in accident-related care among persons who received such prescriptions relative to those who did not remained even after we controlled for prior differences in patterns of health care utilization. Finally, while a greater number of outpatient visits might be related to differences in propensity to seek care, an excess of accident-related hospital admissions among benzodiazepine users can less plausibly be attributed to patients' discretion.

The foremost limitation of our study is that we could observe neither benzodiazepine use nor the occurrence of accidental injuries directly. We inferred the former from the presence in patients' records of pharmacy claims for these agents. We inferred the latter from the presence of nondrug claims with physician-recorded ICD-9-CM diagnosis codes indicating injury as the reason for rendering care, and the assumption that such injuries were accidental and not self-inflicted. To the extent that plan members may have received prescription drugs from other sources, we may have incorrectly identified some persons as nonusers who in fact used benzodiazepines. While it is also possible that some study subjects were treated for injuries without corresponding claims being submitted to the plan, we believe that this is unlikely. Also, there is no apparent reason why such out-of-plan care should be more frequent among persons who were prescribed benzodiazepines than among those who were not, or vice versa.

A more likely problem deriving from the use of claims data is that some of the accident-related care that we observed in the six-month period following initiation of therapy may in fact have been for follow-up treatment of injuries that occurred prior to therapy. The benzodiazepine user group, for example, may have included persons with pre-existing back strain, for which benzodiazepines are sometimes prescribed. Similarly, the benzodiazepine non-user group, which was selected from among all persons with nonbenzodiazepine drug claims, may have included patients who were prescribed analgesics for control of prior accident-related pain. We note, though, that when we compared rates of accident-related encounters among persons who had no such care during the three-month period prior to therapy, our results were largely unchanged.

There are important limitations inherent in any study that uses claims data. We note, however, that data bases of the type we have employed are being used increasingly to study outcomes of care: health care claims data, for example, recently have been used to monitor the incidence of birth defects among patients receiving the acne medication, isotretinoin,[13] and to estimate the risks of mortality and reoperation following prostatectomy.[14]

Despite possible limitations, we believe that our study has several important implications. Our findings suggest that the use of benzodiazepine tranquilizers may increase the risk of accidental injury. They therefore corroborate those of previous studies. The conditions for which these agents are prescribed, however, need and deserve treatment. Accordingly, patients should be advised of the potential risks associated with the use of these agents. Also, when appropriate, alternative therapies might be considered, particularly for persons who must operate machinery or a motor vehicle, those who would use alcohol, and those who may require an extended period of therapy.

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A Case-Control Study of the Effectiveness of Bicycle Safety Helmets in Preventing Facial Injury

Diane C. Thompson, MS, Robert S. Thompson, MD, Frederick P. Rivara, MD, MPH, and Marsha E. Wolf, PhD

Abstract: In a case-control study we sought to assess the potential effectiveness of helmets in preventing facial injuries. Our study included 212 bicyclists with facial injuries and 319 controls with injuries to other body areas, who were treated in emergency rooms of five Seattle area hospitals over a one-year period. Using regression analyses to control for age, sex, education and income, accident severity, and cycling experience we found no definite effect of helmets on the risk of serious facial injury (odds ratio 0.81; 95 percent confidence interval = 0.45, 1.5), but protection against serious injuries to the upper face (odds ratio 0.27; 95% CI = 0.1, 0.8). No protection was found against serious injuries to the lower face. The independent effect of helmet use on facial injury was difficult to isolate due to the association of head and facial injuries.

Our results suggest that bicycle helmets as presently designed may have some protective effect against serious upper facial injuries. (*Am J Public Health* 1990;80:1471–1474.)

Introduction

Recently we showed that bicycle helmets reduce the risk of head injury by 85 percent.[1] Facial injuries in the population we studied occur at nearly identical rates (43 per 100,000) as head injuries (45 per 100,000) in bicyclists and can be serious. Bicycle crashes were found to account for 7.1 percent of all facial fractures treated in a large Swedish University hospital over a three-year period.[2]

There are no reported studies on the effectiveness of bicycle helmets in altering the risk of facial injury. Several reports suggest facial injuries are decreased in helmeted motorcyclists and moped riders,[3–5] but motorcycle and moped helmets generally have face pieces whereas traditional bicycle helmets have no face pieces.

The present case control study describes facial injuries resulting from bicycle crashes and evaluates the effect of bicycle helmets on the risk of these injuries.

Methods

Data for the present study were collected as part of a case control study of head injuries resulting from bicycle crashes. The methods have been described in detail elsewhere.[1] In brief, the study was conducted by emergency room surveillance at five major Seattle area hospitals: the Central and Eastside Hospitals of Group Health Cooperative (GHC) of Puget Sound, a large staff model health maintenance organization; University Hospital of the University of Washington; Harborview Medical Center, a regional Level I trauma center; and Overlake Hospital, a community hospital caring for a large proportion of the trauma to the east of Seattle. The records of the medical examiner's office were examined for the study period to ascertain any cyclist deaths occurring at the scene of a crash.

Definitions

Cases were defined as bicyclists who had suffered a facial injury in a bicycle crash and who sought treatment at the emergency room of one of the five study hospitals during the one-year study period December 1, 1986 through November 30, 1987. Facial injury was defined as injury to the area of the face beginning with the eyebrows and extending to the chin, mouth, and teeth. Injuries to the forehead and skull were classified as head injuries and those without concurrent facial injuries were excluded from the facial injury case group.

To address possible overascertainment of minor facial injuries among those individuals who seek care for more serious concomitant injuries, such as head injuries, we used only cases with more serious facial injury, defined as lacerations, fractures of the facial bones, and fractures of the teeth.[6,7]

Patients with serious facial injury were divided into two groups: the upper facial region included the midface, nose and eye/orbit area, the region where helmets would be expected to have the greatest protective effects; the lower facial region included the mouth and chin.

The case groups formed by the above definitions contained bicyclists who had injuries to both the head and facial regions and those who had face injury without concomitant head injuries. It is difficult to separate the possible effect of an intervention (helmet) on different injuries particularly when there is close anatomical association (head and face). We felt the most appropriate definition of a facial injury case group would include all crashing cyclists who sustained a facial injury regardless of the presence of injuries to other areas of the body.

The main comparison group (N = 319) was composed of bicyclists who sought care at the same five emergency rooms for bicycle related injuries other than to the face or head. Head injuries were excluded from the control group, since a cause and effect relationship between bicycle safety helmet use and head injury exists.

A population-based control group (N = 558) consisting of GHC member bicyclists who had experienced a crash in the last year was employed in ancillary analyses. Those under 15 years of age in this control group had significantly higher helmet usage rates at the time of the cycle crash (21.1 percent)

Address reprint requests to Diane C. Thompson, MS, Research Coordinator, Harborview Injury Prevention and Research Center, 325 Ninth Avenue, ZXC-10, Seattle, WA 98104. Dr. Rivara is Director of the Harborview Injury Prevention and Research Center, and is also Professor of Pediatrics and Adjunct Professor of Epidemiology at the University of Washington School of Medicine; Dr. Wolf is Epidemiologist with the Center, and Assistant Professor of Epidemiology at the University; Dr. Robert Thompson is Senior Director of Preventive Care Research, Group Health Cooperative of Puget Sound. This paper, submitted to the Journal November 16, 1989, was revised and accepted for publication April 11, 1990.

as compared to the same age group from the emergency room controls (5.9 percent).[1]

Data Collection

As described elsewhere[1] emergency room cases and controls were identified by weekly surveillance of emergency room logs. All cases and controls were mailed an introductory letter and study consent form explaining the purpose of the study. Detailed questionnaires described earlier[1] were sent to each person with telephone follow-up of nonrespondents. Medical data on injuries was obtained by abstracting the emergency room records.

Statistical Analysis

The statistical package SASPC was used to evaluate descriptive information and prepare crude risk estimates. Unconditional logistic regression techniques were used to estimate the relative risk of facial injury associated with helmet use and to control for multiple confounding factors.[8] The Egret software package was used to carry out logistic regression. The Mantel Haenszel procedure was used to compute age stratified odds ratios in instances where cell sizes were small or included zero.[9]

Results

Seven hundred seventy-six people were treated in the five hospital emergency rooms for bicycling injuries during the study period. Two hundred fifty-six individuals had facial injuries and were potential cases; 368 had injuries involving other parts of the body and were potential emergency room controls. An additional 151 emergency room controls with head injuries were excluded from the analyses. Overall response rate to our questionnaire was 85.0 percent (82.8 percent of cases and 86.7 percent of controls) resulting in 212 cases and 319 emergency room controls available for analysis.

The type and location of the injuries are presented in Table 1. Of the 212 individuals with facial injuries, 136 (64.2 percent) had injuries to one area of the face, 53 (25 percent) had injuries to two areas of the face, 18 (8.5 percent) had injuries to three areas, and 2.3 percent had injuries to four or five areas of the face.

There were three deaths in the study, all of whom were adults with severe head injuries in addition to facial injuries. No bicyclists died at the crash scene.

Demographic Characteristics and Helmet Use of Cases and Controls

As shown in Table 2, cases were younger than controls. There were no important differences between case and control groups by sex, education of head of household, or household income. The distribution of helmet use at time of crash differed between cases and controls. Individuals with any type of facial injury were less likely to have been using helmets at the time of the crash than the control group. There was substantially lower helmet usage among the youngest cyclists. Helmet ownership rates followed a pattern similar to usage rates. Approximately 90 percent of the helmets were of the hardshell variety.

TABLE 1—Type and Location of Facial Injuries Experienced (N = 212 individuals)

	Number and Type of Injury		
Location	Abrasion/Contusion	Laceration	Fracture
Eye/Orbit	23	8	4 (orbit)
Nose	20	12	4
Midface	72	18	7 (zygoma)
Mouth	12	20	33 (teeth & maxilla)
Chin	29	52	4 (mandible)

Notes: Injuries to one area/location of the face are not mutually exclusive of injuries to other areas.

Types of injuries at each location are mutually exclusive.

Crash Characteristics of Cases and Control

Table 3 portrays various aspects of crash severity. Cases were more likely to have hit a stationary obstacle than were controls. Damage to the bicycle was somewhat more frequent for cases. Collisions with motor vehicles, self-reported bicycle speed at the time of crash, and the type of surface impacted was similar for both cases and controls. Cycling experience (kilometers ridden per week, data not shown) was significantly greater in control group.

Risk of Facial Injury Adjusted for Covariates

The focus of our study was to examine the association between bicycle safety helmet use and a set of successively more restrictive and biologically plausible outcomes: all facial injuries, serious facial injuries, and serious upper facial injuries.

The following variables were evaluated using logistic regression analyses: age, sex, cycling experience (kilometers per week cycled), severity (cause of crash, damage to cycle), education, income, and hospital. Only age, as a covariate, accounted for marked changes in computed odds ratios; therefore, we chose to use the simplest model adjusting only for the effects of age. There appeared to be no definite association between safety helmet use and all facial injuries (odds ratio 0.69; 95% CI = 0.41, 1.1) or serious facial injuries (odds ratio 0.81; 95% CI = 0.45, 1.5). When cases with only serious upper facial injuries were examined, a protective effect emerged (odds ratio 0.27; 95% CI = 0.1, 0.8).

Independent Effect of Helmet Use

The analysis of the association between bicycle safety helmet use and facial injury is difficult because some of the association demonstrated between helmet use and facial injury risk is due to the additional association of facial injury and head injury.

Ninety-eight (46.2 percent) of the 212 facial injury cases had a concurrent head injury; 48 (37.8 percent) of the 127 cases with serious facial injury (lacerations, fractured bones or teeth) had a concurrent head injury. The serious facial injury group included a total of 44 individuals with injuries to the upper facial region and 99 individuals with injuries to the lower face; 16 of those individuals had injuries to both regions of the face. Concurrent head injury was present in 27 (61.4 percent) individuals with serious upper facial injury, and in 31 (31.3 percent) of those with serious lower facial injuries.

A series of ancillary analyses were undertaken to attempt to isolate the effect of helmets on facial injuries. The age-adjusted odds ratio for all cases with concurrent head and face injuries (N = 98) was 0.4 (95% CI = 0.18, 0.83). However, when only serious lower facial injuries were examined, the age-adjusted odds ratio was 1.0 (95% CI = 0.6, 2.0) for cases with facial injury only (N = 114) 1.1, (95% CI = 0.57, 1.97).

Finally, we performed analyses utilizing the population control group of GHC cyclists experiencing a crash in the prior year to address the question of protective effect in crashing cyclists in general. Protective effects were found for

TABLE 2—Demographic Characteristics and Helmet Use of Cases and Controls

Characteristics	All Facial Injury Cases (N = 212) N	%	Serious Facial Injuries (N = 127) N	%	Serious Upper Facial Injuries (N = 44) N	%	Emergency Room Controls (N = 319) N	%
Sex								
Male	155	(73.1)	89	(70.1)	30	(68.2)	224	(70.2)
Female	57	(26.9)	38	(29.9)	14	(31.8)	95	(29.8)
Age (years)								
2–14	138	(65.1)	81	(63.8)	23	(52.3)	126	(39.5)
15–24	24	(11.3)	14	(11.0)	6	(13.6)	86	(27.0)
≥25	50	(23.6)	32	(25.2)	15	(34.1)	107	(33.5)
Education of Head of Household								
High School or less	36	(17.0)	17	(13.4)	5	(11.4)	39	(12.3)
College	116	(54.7)	72	(56.7)	25	(56.8)	168	(53.0)
Graduate school	50	(23.4)	31	(24.4)	13	(29.6)	91	(28.5)
Refusals & unknown	10	(4.7)	7	(5.5)	1	(2.3)	21	(6.6)
Household Income								
<$15,000	32	(15.1)	15	(11.8)	9	(20.5)	56	(17.6)
$15–$35,000	65	(30.7)	33	(26.0)	10	(22.7)	79	(24.8)
≥$35,000	89	(42.0)	62	(48.8)	20	(45.5)	138	(43.3)
Refusals & unknown	26	(12.3)	17	(13.4)	5	(11.4)	46	(14.4)
Wearing Helmet at Time of Crash	30	(14.2)	21	(16.5)	4	(9.1)	83	(26.0)

TABLE 3—Characteristics Contributing to Crash Severity

	Cases (N = 212) N	%	Serious Facial Injuries (N = 127) N	%	Upper Serious Facial Injuries (N = 44) N	%	ER Controls (N = 319) N	%
Cause of Crash								
Contact with moving motor vehicle	30	(14.2)	15	(11.8)	7	(15.9)	51	(16.0)
Contact with other moving objects[a]	16	(7.6)	9	(7.1)	2	(4.6)	40	(12.5)
Contact with stationary obstacle[b]	70	(33.0)	45	(35.4)	16	(36.4)	74	(23.2)
Bicycle malfunction[c]								
Riding unsafe bicycle	21	(9.9)	12	(9.4)	2	(4.6)	27	(8.5)
Falls and other	75	(35.3)	46	(36.3)	4	(9.1)	127	(39.8)
Nature of Surface Involved								
Paved (concrete, asphalt)	188	(88.7)	111	(87.4)	37	(84.1)	268	(84.0)
Gravel	14	(6.0)	10	(7.9)	3	(6.8)	21	(6.6)
Dirt/grass	10	(4.7)	6	(4.7)	4	(9.1)	30	(9.4)
Self-Reported Speed at Time of Crash								
Slow (<5 MPH)	70	(33.0)	42	(33.1)	16	(36.4)	96	(30.1)
Moderate (5–15 MPH)	73	(34.4)	46	(36.2)	18	(40.9)	123	(38.6)
Fast (>15 MPH)	48	(22.6)	28	(22.1)	10	(22.7)	87	(27.3)
Unknown	21	(9.9)	11	(8.7)	0	(0)		
Damage to Bicycle								
No	99	(46.7)	63	(49.6)	24	(54.6)	176	(55.2)
Yes	113	(53.3)	64	(50.4)	20	(45.6)	143	(44.8)

a) Bicycles, cars, pedestrians, animals
b) Parked cars, bumps, curbs
c) Chain fell off, flat tire, handlebars came off, no seat

serious upper facial injury; the age-adjusted odds ratio was 0.14 (95% CI = 0.04, 0.47). When facial injuries only were used as the case group, the odds ratio was 0.5, (95% CI = 0.3, 0.9).

Discussion

The results of this study are consistent in suggesting a protective effect of bicycle safety helmets on serious injuries of the upper facial region (lacerations and fractures of the midface, nose, eye/orbit). This conclusion has biological plausibility. The precise magnitude of the protective effect cannot be estimated, since it was not possible to completely disentangle protective effects against facial injury from the effects of head injury occurring concurrently in the same individual.

The significant association of head trauma and facial injury requires careful case and control definition. Head-injured patients were excluded from the emergency room control group because the head injury is directly related to the exposure (helmet use).[1] Their inclusion would have resulted in an underestimation of the protective effect of helmets on facial injuries by selecting for unhelmeted controls. This would be analogous to excluding individuals with lung cancer from the control group in a study of the association of smoking with heart disease.

The most representative facial injury case group is composed of all cyclists with facial injuries including those with injuries to the head or other portions of the body. Equal ascertainment of facial injuries among injured cyclists is important. The presence of a concurrent head injury increases the likelihood that an injured individual seeks medical attention and thus becomes a case. Since head injury is so strongly associated with the lack of helmet use, such a bias can result in an overestimation of the protective effect of helmets by potentially increasing the number of unhelmeted cases. Ascertainment bias was minimized by defining a case group with serious facial injury, i.e. injury serious enough to necessitate medical attention independently of the presence of a head injury. This analysis suggests no association between helmet use and the risk of facial injuries.

We attempted in the analysis to determine the independent effect of bicycle safety helmets on facial injuries. Although there was no effect of helmets on facial injury alone (cases with concurrent head injury excluded) when the emergency room controls were used as comparison, use of the same case group with the population-based control group suggested some protective effect of helmet use (odds ratio 0.5; 95% CI = 0.3, 0.9). This is a more conservative analysis since individuals with head injury and thus low chance of helmet use were excluded. If a helmet effect is present, one would expect to show this effect using the population controls as the comparison group. This is a less seriously injured group of cyclists and has a higher rate of helmet use than the emergency room controls. The demonstration of some protective effect with facial injuries only is important, since it indicates effect on facial injury separate from head injury effects.

Current bike helmets do not have a face bar; consequently one would expect little or no protection against injuries to the lower face. The negative results for serious lower facial injury lend credence to the positive protective effect found for serious upper facial injury.

The strong association between helmet use and decreased risk for serious upper facial injuries is likely explained by a combination of factors: the association of upper facial injuries with head injuries, (61.4 percent of those with serious upper facial injuries) and an actual protective effect of helmets for upper facial injury.

A potential limitation of this study is that unmeasured factors not taken into account in our analysis may have placed controls at lesser risk for striking their faces at the time of the crash.

In conclusion, our study suggests that bicycle helmets as presently designed have little or no effect on the overall risk of facial injury, but appear to provide some protection for serious upper facial injuries. The magnitude of this effect is not precisely quantifiable. This study also illustrates the methodologic problems of isolating exposure or intervention effects on one injury type when multiple simultaneous injuries result from a single event. This problem is inherent to injury research, but much less likely in cancer or cardiovascular investigations.

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Association of Hepatitis B Surface Antigen and Core Antibody with Acquisition and Manifestations of Human Immunodeficiency Virus Type 1 (HIV-1) Infection

RACHEL E. SOLOMON, MHS, MARK VANRADEN, MA, RICHARD A. KASLOW, MD, MPH, DAVID LYTER, MD, BARBARA VISSCHER, MD, HOMAYON FARZADEGAN, PhD, AND JOHN PHAIR, MD

Abstract: We examined the associations between seropositivity for hepatitis B virus (HBV) with the presence or development of antibodies to human immunodeficiency virus (HIV-1) and with HIV-1 induced T-helper lymphocyte deficiency or acquired immunodeficiency syndrome (AIDS). Serologic data on HBV and HIV-1, cytometric enumeration of CD4$^+$ lymphocytes, clinical events (AIDS by Centers for Disease Control criteria) and hepatitis B vaccination histories were available on 4,498 homosexual participants in the Multicenter AIDS Cohort Study. Men were classified as to previous infection with HBV and prevalent or incident infection with HIV-1. Although there was an association between seropositivity for HBV infection and HIV-1 infection at enrollment (odds ratios anti-HB$_c$ 2.6; HB$_s$Ag 4.2), the relation between HBV seropositivity and subsequent seroconversion to HIV-1 was weaker (odds ratios 1.3 and 1.6). HIV-1 seroconversion was also associated with a history of certain other sexually transmitted diseases, but predisposing sexual practices did not account for the association between HBV and HIV-1 infection. Seropositivity for HBV infection at entry was not related to initially low or more rapid subsequent decline in T-helper lymphocyte counts and was not associated with an increased incidence of AIDS during 2.5 years of follow-up. History of vaccination against HBV did not appear to decrease susceptibility to HIV-1 infection or to subsequent progression of immunodeficiency. We conclude that prior HBV infection is unlikely to be specifically associated with acquisition of HIV-1 infection and is unrelated to more rapid progression of HIV-1-induced immunodeficiency. (*Am J Public Health* 1990; 80:1475–1478.)

Introduction

Approximately 300,000 new cases of hepatitis B virus (HBV) infection occur every year in the United States alone.[1] Before the spread of acquired immunodeficiency syndrome (AIDS) resulted in altered patterns of sexual behavior, an estimated 60 percent of homosexual males demonstrated serologic evidence of previous hepatitis B infection as measured by antibody to hepatitis B core antigen (anti-HB$_c$), and approximately 5 percent were chronic carriers of hepatitis B surface antigen (HB$_s$Ag).[2] Homosexual males were at 10 to 15 times greater risk of prior HBV infection than men in the general population.[2] Consequently, even before human immunodeficiency virus, type 1 (HIV-1) was established as the cause of AIDS there was considerable interest in a possible direct or cofactorial role for HBV in the AIDS epidemic.[3–5]

In most epidemiologic studies, homosexual men have experienced high rates of infection with both HBV and HIV-1.[6–9] Detailed cross-sectional analysis of men who enrolled in the Multicenter AIDS Cohort Study (MACS), a prospective study of HIV-1 infection in homosexual men, demonstrated that the close association between detectable anti-HB$_c$ or HB$_s$Ag and seropositivity for HIV-1 infection was highly significant and independent of sexual practices commonly predisposing to both infections.[10] Observations of this kind, coupled with the recent description of an effect of the hepatitis B virus protein X on HIV-1 replication in vitro,[11,12] have encouraged speculation about concurrent HBV as a codeterminant for infection or disease with HIV-1. We report here an analysis of possible relations of prevalent HBV infection (anti-HB$_c$ or HB$_s$Ag) to the acquisition and to manifestations of HIV-1 infection in homosexual men.

Methods

The MACS enrolled 4,954 homosexual males without AIDS as defined by CDC criteria[13] in four metropolitan US areas[9,10] from April 1984 through April 1985. The study centers are located in Baltimore, Maryland (Johns Hopkins University), Chicago, Illinois (Howard Brown Memorial Clinic/Northwestern University), Los Angeles, California (University of California), and Pittsburgh, Pennsylvania (University of Pittsburgh). At both entry and follow-up visits interviewers administered a standardized questionnaire, through which we obtained a demographic profile, past medical history, drug and medication history, history of hepatitis B vaccination, psychological assessment, and sexual behavior data.

Serum specimens taken from all men at entry were examined for antibodies to HIV-1, for HB$_s$Ag, and for anti-HB$_c$; at each six-month follow-up visit participants were tested for antibodies to HIV-1. Testing for anti-HIV-1 was performed by enzyme-linked immunoassay (EIA) using licensed, commercially available reagents (DuPont and Genetic Systems), and positive results on the EIA were confirmed by immunoblot assays. Standard EIA or radioimmunoassay techniques using commercially available kits (Abbott Laboratories, Inc.) were used to test for prevalent HBV infection (defined as detectable HB$_s$Ag or anti-HB$_c$). The ordinarily small number of men who may have had antibodies to HB$_s$Ag only were thereby classified as HBV-seronegative. Lymphocytes were phenotyped using standard flow cytometric techniques, as previously described for the MACS,[9] at each of the first three visits for all men and at all subsequent visits for HIV-1-seropositive men, as well as for a subset of HIV-1-seronegative men.

From the National Institute of Allergy and Infectious Diseases (Solomon, VanRaden, Kaslow); the University of Pittsburgh Graduate School of Public Health (Lyter); UCLA Department of Epidemiology (Visscher); Department of Epidemiology, Johns Hopkins School of Hygiene and Public Health, (Farzadegan); Howard Brown Memorial Clinic/Northwestern University Medical School, Department of Medicine (Phair). Address reprint requests to Richard A. Kaslow, MD, MPH, Epidemiology and Biometry Branch, DMID, NIAID, Westwood 739, National Institutes of Health, Bethesda, MD 20892. This paper, submitted to the Journal December 18, 1989, was revised and accepted for publication May 21, 1990.

In our analysis of the effects of hepatitis B infection on HIV-1 infection and accompanying immunodeficiency, we excluded 287 men whose initial serologic status (for HBV or HIV-1) or hepatitis B vaccination status was unknown. Of the remaining 4,667 participants, 531 reported having received hepatitis B vaccine prior to entry, and 362 of those vaccinated were known to be serologically negative for HB$_s$Ag and anti-HB$_c$. These men were analyzed separately because of the diminished likelihood of past or current HBV infection.

To examine the possible relation between prior HBV infection and the degree of cellular immunodeficiency, we compared HBV-infected and -uninfected HIV-1-seropositive men for differences in their numbers of CD4$^+$ and CD8$^+$ lymphocytes at entry and differences in the mean changes in those numbers during 2.5 years of follow-up. Slopes of decline in CD4$^+$ cells were calculated by the least squares method, and differences between slopes for the HBV-infected and -uninfected men were compared with a standard t or a Wilcoxon rank sum test. In addition, a life table analysis of the first 2.5 years of follow-up in HIV-1-seropositive MACS participants was performed to examine the effect of prior HBV infection on the incidence of AIDS.

Associations between outcomes and potential risk factors were individually estimated with odds ratios (OR). Life tables and associated log rank tests were calculated by standard computerized statistical packages (SAS Proc Survfit and Proc Survtest). For cross-sectional multivariate analyses, multiple logistic regression techniques were employed using a standard package (BMDPLR), and adjusted ORs were computed from the coefficients derived from the regression model. For prospective analyses, time-dependent Cox proportional hazards models were applied using a standard package (SAS Proc PHGLM).

Results

The 2,826 MACS participants who were unvaccinated against HBV and had serologic evidence of prior or current HBV infection at entry differed from their 1,310 HBV-seronegative counterparts in HIV-1 seroprevalence at entry and in rates of HIV-1 seroconversion during 6 and 24 months of follow-up (Table 1). For further analysis of the relation between hepatitis B and HIV-1 infections present at baseline, multiple logistic regression was used to adjust for a number of known or suspected baseline correlates of anti-HIV-1 (Table 1). The presence of HB$_s$Ag and/or anti-HB$_c$ remained independently correlated with the presence of HIV-1 at entry (adjusted OR: presence of HB$_s$Ag ± anti-HB$_c$, 4.2; presence of anti-HB$_c$ only, 2.6) (Table 1). As previously reported from the MACS,[10] other sexually transmitted infections such as gonorrhea, anogenital herpes, syphilis, cytomegalovirus infection and others, taken singly or in combination, also showed strong relations (ORs = 1.7–4.1) to baseline HIV-1 seropositivity (data not shown).

In the analysis of the relation between baseline HBV seropositivity and subsequent HIV-1 seroconversion among the 2,421 unvaccinated HIV-1-seronegative men, after adjustment for other potentially related variables, the presence of HB$_s$Ag and of anti-HB$_c$ at entry was less strongly but still related to HIV-1 seroconversion in 83 men who developed HIV-1 antibodies during the six-month follow-up and in 181 who developed HIV-1 antibodies during the 30-month follow-up (Table 1). The analyses also suggested independent relations of magnitude similar to those seen in the cross-sectional analyses (OR = 1.4–1.9) between HIV-1 seroconversion and a history of one or more sexually transmitted infections other than HBV.

To determine the possible role of previous HBV infection in the progression of quantitative T-lymphocyte deficiency, we analyzed T-lymphocyte phenotypes in unvaccinated HIV-1-seropositive participants as follows: 1,348 men with six months, 1,131 men with 18 months, and 929 men with 30 months of follow-up data. The presence of HB$_s$Ag or

TABLE 1—Baseline Prevalence and 6-Month and 24-Month Incidence Rates of HIV-1 Infection According to Baseline HBV Serologic Status in 4,136 Homosexual Men

Baseline HIV-1 Serologic Status (N = 4136)

Baseline HBV Status	Number Studied[*]	HIV-1 Seropositive %	Relative Risk[+] of Baseline HIV-1 Seropositivity: Univariate	Relative Risk[+] of Baseline HIV-1 Seropositivity: Multivariate[†]
HB$_s$Ag+, anti-HBc+or−	282	52.1	6.1	4.2 (2.9–5.9)[§]
HB$_s$Ag−, anti-HBc+	2544	48.3	5.3	2.6 (2.1–3.2)
HB$_s$Ag−, anti-HBc−	1310	15.0	1.0	1.0

Follow-up of Initially HIV-1 Seronegative Men (N = 2421)

Baseline HBV Status	Number Studied[*]	HIV-1 Seropositive %[**]		Relative Risk[†][+] of HIV-1 Seroconversion: Univariate		Relative Risk[†][+] of HIV-1 Seroconversion: Multivariate[†]	
		6m	24m	6m	24m	6m	24m
HB$_s$Ag+, anti-HBc+or−	122	4.1	7.6	2.6	1.7	2.2 (0.8–6.5)	1.3 (0.7–2.7)
HB$_s$Ag−, anti-HBc+	1253	4.9	10.4	3.1	2.3	2.2 (1.2–4.1)	1.6 (1.1–2.3)
HB$_s$Ag−, anti-HBc−	1046	1.6	4.7	1.0	1.0	1.0	1.0

[*] Unvaccinated against hepatitis B.

[+] For baseline and 6-month follow-up analyses, relative risk was measured by odds ratio; for 24-month follow-up analyses, relative risk was measured by hazard ratio (see Methods).

[†] Multivariate analysis adjusted for age, race, study center, education, indices of sexual activity (types and frequencies of specific sexual practices, history of other sexually transmitted infections, a serologic test for syphilis and antibodies to cytomegalovirus) and both parenteral and nonparenteral psychoactive drug use.

[§] 95% confidence intervals by Woolf method.

[**] Cumulative proportions by Kaplan-Meier method.

anti-HB$_c$ showed no consistent relation either to mean initial T-helper (CD4$^+$) and T-suppressor (CD8$^+$) cell counts or to any change in counts of these cell types during the follow-up periods. The absence of any effect on mean decline in CD4$^+$ cell counts persisted after adjustment for initial count. During the entire follow-up, among unvaccinated men the mean six-month declines in counts were: all HB$_s$Ag+ = .0391 cells × 10^9/L; HB$_s$Ag−, anti-HB$_c$+ = .0452 cells × 10^9/L; and HB$_s$Ag−, anti-HB$_c$− = .0421 cells × 10^9/L (p = NS). Comparable analyses were performed on T-lymphocyte data available for approximately 18 months from HIV-1-seronegative men, and no relations were found.

During the first 2.5 years of follow-up, AIDS developed in 198 of the 1,532 participants who had HIV-1 antibodies at entry and had not received hepatitis B vaccine. The cumulative 2.5 year life table incidence of AIDS in these men was the same whether serologic evidence of HBV infection was present or absent at entry (Figure 1). In particular, neither men who were positive for HB$_s$Ag at baseline nor men who were negative for HB$_s$Ag but positive for anti-HB$_c$ had a higher incidence of AIDS than men who were negative for both HB$_s$Ag and anti-HB$_c$. The results were unchanged when stratified by number of CD4$^+$ cells at entry.

To assess the possible effects of HBV vaccine on susceptibility to HIV-1 infection, we analyzed HIV-1 seroconversion rates in HIV-1-negative, HB$_s$Ag-negative, anti-HB$_c$-negative vaccine recipients versus serologically identical non-recipients. Among 1,320 such participants who had completed at least six months of follow-up, men with and without a history of vaccination reported similar high-risk sexual practices and seroconverted at similar rates during the next six and the next 24 months. In the 24-month period, life table rate of HIV-1 seroconversion was not substantially different between vaccinees (n = 274, 6.0%) and nonvaccinees (n = 1,046, 4.7%) (95% confidence interval of difference: −1.9, 4.5). The slightly higher seroconversion rate in vaccinees is consistent with their slightly higher frequencies of such indicators of high risk as the numbers of sexual partners in the recent past and the proportion of men who had had at least one sexually transmitted infection (data not shown).

At entry, HIV-1-seropositive men with a history of HBV vaccination did not demonstrate a low CD4$^+$ cell count more frequently than serologically identical unvaccinated men. During the 2.5 years of follow-up, the vaccinated men showed a slightly more rapid decline in their CD4$^+$ counts. The difference in mean slopes between vaccinated men (−.075 cells × 10^9/L per six months) and unvaccinated men (−.042 cells × 10^9/L per six months) fluctuated with the statistic examined (p = .044 by t test, p = .017 by the nonparametric Wilcoxon test). During the first 2.5 years of follow-up of initially HIV-1 seropositive men, the life table incidence of AIDS was about 15 percent in those who were HBV-seropositive (Figure 1). In the same interval the incidence was about 20 percent in those who were HBV-seronegative with or without a history of HBV vaccination.

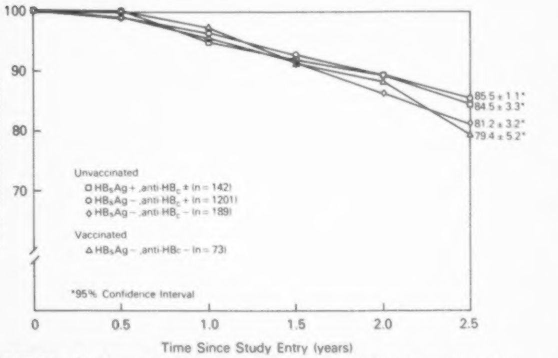

FIGURE 1—Incidence of AIDS among 1605 Initially HIV-1 Seropositive Men, According to Hepatitis B Virus Serologic Status at Baseline

Discussion

Recent evidence that HBV can infect lymphocytes and produce a protein capable of activating HIV-1[11,12] has renewed interest in the early hypothesis that HBV influences susceptibility to or progression of infection with HIV-1.[14,15] Previous studies including our own[10] have demonstrated strong associations between the presence of HIV-1 antibodies at study entry and previous or persistent infection with several sexually transmitted infections, including HBV. Our current investigation confirmed these cross-sectional observations; however, the more powerful longitudinal analysis has provided only meager evidence that previous HBV infection is truly a cofactor for acquisition of infection with HIV-1. We found that HBV infection at baseline was weakly correlated with HIV-1 seroconversion during the next six and the next 24 months. This relation could imply that past or current HBV infection facilitates HIV-1 infection. Since the relation to seroconversion in both time intervals occurred among individuals positive only for anti-HB$_c$, if there is a true relation, it does not appear to depend closely on the presence of circulating HB$_s$Ag. Although the data could be interpreted as indicative of a causal relation, considering the weakness of the association, we believe that other explanations must be considered equally plausible. One alternative explanation would be that HBV is but one of several sexually transmitted infections nonspecifically capable of facilitating establishment of HIV-1 infection, for example by "priming" T-lymphocytes for penetration or viral replication. Our analyses demonstrated independent relations of similar magnitude between several other sexually transmitted diseases and HIV-1 seroconversion. However, such multiple interrelations could be taken as evidence against any causal contribution of HBV.

To allow for possible temporal variations in the appearance of HIV-1 antibody, we further adjusted for the timing of HIV seroconversion by excluding men who seroconverted early in the follow-up period (within nine months of the baseline evaluation). This did not alter the association between anti-HB$_c$ seropositivity and HIV-1 seroconversion.

Although HBV may persist among persons expressing anti-HB$_c$ in the absence of HB$_s$Ag, a causal relation might have been more readily demonstrable among HB$_s$Ag-positive men. There was no clear evidence of such a causal relation to HIV-1 seroconversion. The absence of this relation was not likely due to the small number of men in this group: the power to detect a difference of two-fold or greater between them and all HB$_s$Ag-negative men was 80 percent.

Our observation that the HIV-1 seroconversion rates were comparable in vaccinated and in unvaccinated seronegative men argues against a protective effect of vaccine and against such a causal role for HBV. A few participants may have erroneously reported HBV vaccination or overestimated its completeness, but it is highly unlikely that misclassified vaccination experience explains these findings.

Our investigation provided no evidence that infection with HBV hastens the appearance of the major immunologic or clinical manifestations of HIV-1 infection. We found no firm associations between positive HBV serology and either altered T-cell phenotype distribution at the time HIV-1 seropositive men entered the study or declining numbers of T-helper cells during the next 24 months. Somewhat divergent results have been reported from the few other studies of the effects of HBV infection on T-cells[16–19] or on clinically significant immunodeficiency[20] in HIV-1-seropositive men. The divergence of findings could be due to the small size of the other studies or to differences in clinical settings, types of patients (i.e. heterosexual vs homosexual) or relative stages of HIV-1 infection. The apparent protection from progressive immunodeficiency in HIV-1-seropositive homosexual male HB$_s$Ag carries in San Francisco[20] was of borderline significance, and the analysis used in that study may not have fully accounted for age, the timing of HIV-1 infection or other factors among the HB$_s$Ag carriers. Our investigation also failed to demonstrate any association, before or after stratification by initial CD4$^+$ cell count, between HBV serologic status and subsequent development of AIDS.

A history of HBV vaccination was not related to altered T-helper lymphocyte counts at entry and probably not related to the slightly more rapid decline in counts subsequently. Our findings, like those of Stevens, *et al*,[20] and others, also suggested that hepatitis B vaccination neither enhances nor retards the development of AIDS. However, our ability to detect an effect of vaccination was limited by the number of vaccinees available for study.

In summary, this prospective investigation of a large number of homosexual men has provided insufficient epidemiologic evidence for and some evidence against the idea that prior or persistent HBV infection facilitates the acquisition of infection with HIV-1, no indication that past or concurrent HBV infection alters progression of HIV-1-induced immunodeficiency, and no suggestion that HBV vaccination in any way protects against either acquisition or progression of HIV-1 infection.

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The Multicenter AIDS Cohort Study Group: Baltimore, Maryland: Johns Hopkins University School of Hygiene and Public Health: Alfred J. Saah, MD, Principal Investigator; Robin Fox, MS; Joseph Margolick, MD; Homayoon Farzadegan, PhD

Chicago, Illinois: Howard Brown Memorial Clinic-Northwestern University Medical School; John P. Phair, MD, Principal Investigator; Joan S. Chmiel, PhD; Steven Wolinsky, MD; Kathleen Sheridan, PhD; Bruce Cohen, MD

Los Angeles, California: University of California Schools of Public Health and Medicine, Los Angeles: Roger Detels, MD, MS, Principal Investigator; Barbara R. Visscher, MD, DrPH; John L. Fahey, MD, MS; Janis V. Giorgi, PhD; Jan Dudley, MPH

Pittsburgh, Pennsylvania: University of Pittsburgh Graduate School of Public Health: Charles R. Rinaldo, Jr, PhD, Principal Investigator; Lawrence A. Kingsley, DrPH; Monto Ho, MD; David W. Lyter, MD

Bethesda, Maryland: National Institutes of Health, National Institute of Allergy and Infectious Diseases: Sten Vermund, MD; Richard A. Kaslow, MD; Mark J. VanRaden, MA; National Cancer Institute: Iris Obrams, MD

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Death in Prison: Changing Mortality Patterns among Male Prisoners in Maryland, 1979–87

MARCEL E. SALIVE, MD, MPH, GORDON S. SMITH, MB, CHB, MPH, AND T. FORDHAM BREWER, MD, MPH

Abstract: The leading causes of death (rate per 100,000 prisoner-years) in Maryland state prisons for the period 1979–87 were circulatory system disease (59), suicide (40), and homicide and legal intervention (30). Acquired immunodeficiency syndrome (AIDS) became the leading cause of death in 1987. Homicides declined after 1980; drug overdose deaths peaked in 1981 and later disappeared. Male inmates have 39 percent lower all-cause death rates than the general population of Maryland after adjustment for age and race. (*Am J Public Health* 1990; 80:1479–1480.)

Introduction

Growing concern with the quality of prison health care has led several groups, notably the American Public Health Association, to develop standards of care for the incarcerated.[1] Despite the recognition that outcomes, including mortality, are useful indicators of quality of care in hospitals[2] and potentially prisons, prison standards do not require routine monitoring. Of over 500,000 inmates in state and federal prisons, 915 males died in the year ending June 1987.[3] National criminal justice statistics[3,4] merely group by type of death, and do not compute rates, make comparisons, or follow trends. Earlier studies[5,6] suggest that inmates have lower mortality than the general population. Descriptive studies have been reported from jails[5,7–10] but not from state prisons where inmates have longer sentences.

This study was designed to: investigate deaths among inmates in the Maryland prison system, examine trends, and compare inmate mortality with that of the general population.

Methods

All male deaths of Maryland state prisoners between January 1, 1979 and December 31, 1987 were studied. Maryland prisons house those persons convicted and sentenced to at least one year; the average stay is 40 months.[11,12] Females comprise only 4 percent of inmates, and were excluded. The prison management information system provided prisoner characteristics. Causes of death obtained from death certificates were classified by nosologists according to the International Classification of Diseases (ICD), 9th edition.[13] Deaths due to illicit drug overdose had been inconsistently classified; for the study they were grouped as unintentional poisoning (ICD E850–E858).

Annual and cause-specific death rates were calculated using prison census figures.[11,12] Indirect standardization used age- and race-specific death rates from Maryland[12,14–16] to calculate expected deaths and standardized mortality ratios (SMR). The 95 percent confidence interval (CI) of the SMR was computed.[17]

Results

Between 1979 and 1987, 204 male inmates died during 93,445 prisoner-years (crude mortality rate 218 per 100,000 prisoner-years). Of the deaths, 140 were from natural causes, 37 suicides, 16 illicit drug overdoses, 9 other unintentional injuries, and two undetermined (Table 1). Leading causes of death (rate per 100,000 prisoner-years) were circulatory system disease (59), suicide (40), homicide and legal intervention (30), malignant neoplasms (28), and unintentional injuries (25).

All-cause mortality peaked in 1980–81 and has since declined paralleling external causes (Figure 1). The natural-cause death rate remained relatively constant, although individual cause rates varied, especially for circulatory system diseases (data available from author on request). Infectious disease death rates were zero until 1984 when they increased rapidly, largely related to human immunodeficiency virus (HIV) infection, and became the leading cause of death in 1987 (65 per 100,000). Other natural causes of death occurred at steady rates. Illicit drug overdose accounted for a rise in external cause death rate that peaked in 1981 (data available from author on request). Homicide and legal intervention deaths had a slight peak in 1980–81 and decreased thereafter. Only one death was due to motor vehicles.

When compared with the Maryland general population,

TABLE 1—Comparison of Mortality among Male Prisoners (1979–87) with That of Men in the Maryland General Population (1983) for Major Causes of Death

	Number of Deaths			
Cause of Death (ICD Code)	Observed	Expected	SMR*	(95% CI)
Infectious and parasitic (001–139)	12	6.1	197	(114, 366)
Malignant neoplasms (140–208)	26	53.9	48	(33, 73)
Endocrine and metabolic (240–279)	1	4.6	22	(1, 121)
Circulatory system (390–459)	55	76.8	72	(54, 93)
Respiratory system (460–519)	9	8.6	105	(48, 199)
Digestive system (520–579)	5	11.4	44	(14, 103)
Genitourinary system (580–629)	3	3.3	90	(19, 264)
External causes (E800–E999)	88	119.4	74	(59, 91)
Unintentional injuries (E800–E949)	23	48.6	47	(23, 87)
Suicide (E950–959)	37	20.1	184	(130, 253)
Homicide and legal intervention (E960–E978)	28	47.5	59	(39, 85)
Other causes of death (290–239, 280–389, 630–799)	5	38.6	13	(4, 30)
All causes	204	334.2	61	(53, 70)

*Standardized mortality ratio

From the Department of Health Policy and Management Johns Hopkins School of Hygiene and Public Health (Salive, Brewer); the Maryland Division of Correction (Salive, Brewer); and the Injury Prevention Center, JHUSHPH (Smith). Address reprint requests to Marcel E. Salive, MD, Center for Medical Effectiveness Research, Agency for Health Care Policy and Research, 5600 Fishers Lane, Rm 18A-19, Rockville, MD 20857. This paper, submitted to the Journal May 3, 1989, was revised and accepted for publication April 9, 1990.

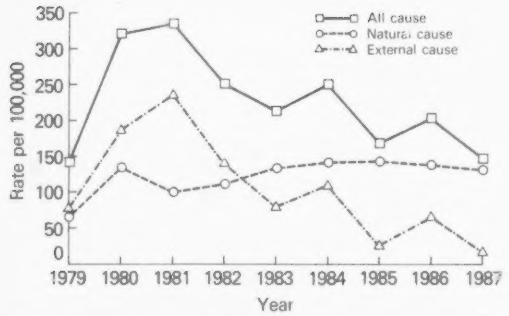

FIGURE 1—All Cause Mortality Rate of Male Inmates over Time, Natural Cause and External Cause Mortality Rates, Maryland State Prison System, 1979–87.

prisoners had lower all-cause mortality for the eight-year period (Table 1) and only suicide and infectious disease deaths were significantly elevated.

Discussion

We examined cause of death in long-term (prison) inmates. Although prisoners had lower all-cause mortality than the general population, significant elevations of suicide and infectious disease deaths were observed. Infectious disease mortality increased due to AIDS, which emerged in 1987 as the leading cause of inmate death. Although fewer than 20 AIDS cases were diagnosed before 1988, the problem will increase since 7 percent of entering male inmates were HIV infected,[18] and HIV transmission has been demonstrated.[19] AIDS prevention and treatment should remain a top correctional health priority. The need for suicide prevention was discussed elsewhere.[12]

Cancer mortality was 46 percent below the comparison population. Circulatory system deaths were also lower, in contrast to prior studies.[6]

External cause mortality overall was 26 percent below that of the general population. Homicide declined steadily and was 41 percent below that of the general Maryland population, although 13 percent of prisoners were convicted murderers. Improved security in the prison, including metal detector installation, is probably responsible for this decline. All injury deaths are potentially preventable and should be investigated.[10]

The epidemic of illicit drug overdose deaths between 1980 and 1982 is probably related to pentazocine (Talwin) use. Six of 16 overdose death certificates reported pentazocine. Pentazocine, although manufactured for oral use, was frequently injected by drug abusers. Its 1983 reformulation added the narcotic antagonist naloxone (inactive when taken orally), reducing overdose deaths nationwide[20,21] and in this prison.

There is a great deal of appropriate concern about poor conditions, riots, murders, outbreaks of illness, and standards of medical care in prisons. Nevertheless, the 39 percent lower Maryland prisoner mortality, similar to that reported in other studies,[5,6] suggests reason for optimism. Decreased mortality may be from a protected environment or lifestyle or a "healthy inmate effect" from selection factors.[5,6] Although data were unavailable from the medical parole program, in our experience premature discharges were rare (none for AIDS) and unlikely to explain lower mortality. Access to medical care may be better in prison than in the general population, which may lead to early detection and treatment of chronic disease. The effect of the availability of correctional officers trained in cardiopulmonary resuscitation needs further study. Limited access to weapons and motor vehicles is likely responsible for low injury mortality, although suicide is elevated. Data limitations preclude an analysis of the development of conditions over time.

The Supreme Court has reiterated the right of prisoners to adequate health care.[22] Any inmate death raises questions about cause, quality of care, and legal responsibility.[5,10] Despite our finding of lower mortality among prisoners, all prisoner death should be investigated to identify remediable health care or safety deficiencies. Improved prison health surveillance is needed for deaths and to determine the effect of prison health services on morbidity.

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Long-Term Impact of Smoking Cessation on the Incidence of Coronary Heart Disease

ANNA N. A. TOSTESON, SCD, MILTON C. WEINSTEIN, PHD, LAWRENCE W. WILLIAMS, MS,
AND LEE GOLDMAN, MD, MPH

Abstract: Using a simulation model of the US male population, we estimated the long-term impact that future smoking cessation programs would have on the distribution and occurrence of coronary heart disease in males ages 35–84. For interventions that reduce the number of smokers by 25 percent in 1990, the number of men free of coronary heart disease is projected to increase by 416,787 (0.7 percent) in 2015, and the age-standardized absolute incidence to decline by 2.3 percent. Incidence rates and absolute incidences are projected to fall in men under age 65, but absolute incidence would rise in men over age 65, in large part because of the increased number of men who were at risk for coronary heart disease because of a reduction in non-coronary smoking-related mortality. These trends were more marked for greater smoking reductions and were generally unaffected in a variety of analyses using alternative assumptions, which considered smoking as a risk factor in the elderly, a lag-time before benefits from smoking cessation were realized and secular declines in smoking prevalence. Subject to the assumptions of our model, we conclude that smoking reductions will markedly reduce coronary heart disease, especially in younger age groups, and that this benefit will be slightly offset by a small increase in absolute incidence in elderly men. (*Am J Public Health* 1990; 80:1481–1486.)

Introduction

Smoking is a major risk factor not only for the development of coronary heart disease, but also for other diseases, such as lung cancer and chronic obstructive pulmonary disease.[1,2] Because these risks are to be reduced in persons who quit smoking,[3–6] a growing literature has investigated the effectiveness, economic consequences, and cost-effectiveness of smoking cessation techniques in clinical practice.[7–11]

However, the impact of smoking cessation programs on one outcome, such as coronary heart disease, may be confounded by its impact on the competing risks of developing other smoking-related diseases. We therefore sought to estimate the long-term effect that smoking intervention programs would have on the primary prevention of coronary heart disease in men who are initially free of heart disease at the time of the intervention. Our analysis focused on male smokers, because the risks of developing coronary heart disease as a function of smoking are more precisely measured for males than for females.[12]

Methods

The Coronary Heart Disease Policy Model, which has been described in full detail elsewhere,[13–15] is a state-transition computer-simulation model consisting of three integrated submodels: the Demographic-Epidemiologic Submodel, the Bridge Submodel, and the Disease History Submodel. The Bridge Submodel characterizes subjects for the first 30 days after they develop coronary heart disease, and the Disease History Submodel considers all events that occur among such persons after that 30-day period. Because our analysis focused on the effect of smoking-reduction programs on primary prevention of coronary heart disease, only the Demographic-Epidemiologic Submodel, which simulates the distribution of coronary heart disease risk factors in a population of 35–84 year old persons who are free of coronary heart disease, and computes the annual projected incidence of coronary heart disease on the basis of the simulated distribution, was used.

Each year, persons age 35 years who remain free of coronary heart disease enter the Demographic-Epidemiologic Submodel of the Coronary Heart Disease Policy Model, and persons who turn age 85 exit from the submodel. To exit from the submodel before age 85, persons must either die of non-coronary heart disease causes or develop coronary heart disease (at which time they enter the Bridge Submodel).

The risk factors used by the Demographic-Epidemiologic Submodel to assess each individual's risk of death and of developing coronary heart disease are age, sex, smoking status (no, yes; if yes, average number of cigarettes per day), diastolic blood pressure (≤94 mmHg, 95–104 mmHg, ≥105 mmHg), relative weight (≤109% of norm, 110–129% of norm, ≥130% of norm), and serum cholesterol (≤249 mg/dl, 250–299 mg/dl, ≥300 mg/dl). Based on the categories for each of these factors, the entire United States population that is free of coronary heart disease and is age 35–84 is divided into 5,400 cells, each with a specific value for each factor. For simplicity purposes, however, the model's outputs are collapsed into 10-year age ranges, thus yielding 540 strata.

Assumptions of the Model

The Demographic-Epidemiologic Submodel was initially constructed using the United States population for 1980 by age and sex[16] and the estimated proportion of persons with no history of coronary heart disease.[17] The distributions of smoking status, diastolic blood pressure, relative weight, and serum cholesterol were taken from the Second Health and Nutrition Examination Survey,[18] and the four risk-factor distributions were, for simplicity purposes, assumed to be independent, conditionally upon age range and sex. This assumption may either over- or underestimate the impact that risk factor changes would have on coronary heart disease. For example, a reduction in smoking could cause a decrease in blood pressure or an increase in weight, and an increase in weight could cause an increase in blood pressure.

The number of US residents who enter the model in each subsequent year after 1981 by turning 35 years of age was

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From the Divisions of Clinical Epidemiology, Cardiovascular, and General Medicine, Brigham and Women's Hospital and the Division of Clinical Epidemiology, Beth Israel Hospital and Harvard Medical School; Department of Health Policy and Management, Harvard School of Public Health. Address reprint requests to Lee Goldman, MD, MPH, Division of Clinical Epidemiology, Department of Medicine, Brigham and Women's Hospital, 75 Francis Street, Boston, MA 02115. This paper, submitted to the Journal December 5, 1989, was revised and accepted for publication May 31, 1990.

estimated from projections of the US Bureau of the Census.[19,20] Risk factor changes with age were estimated from population based cross-sectional data.[18]

For each of the 540 cells, logistic regression coefficients for all-cause mortality and for developing coronary heart disease were based on data from the Framingham Heart Study.[12] The age- and sex-specific coefficients that were used for all risk factors in the Demographic-Epidemiologic submodel were obtained from the Framingham Heart Study's 30-year follow-up, and define the relative odds of death and heart disease for persons with the risk factor in comparison to a baseline group. These raw coefficients were smoothed by age for each sex by means of a weighted least squares regression of the Framingham coefficients on the mid-point of each 10-year age interval, where the weights used were the inverse of the coefficient's standard errors for coronary heart disease incidence. This method gave more weight to the more stable (i.e. smaller standard error) coefficients. The model's coefficients and standard errors that were obtained using this method are compared with the actual Framingham coefficients in Table 1. Note that for persons ≥65 years the smoothed coefficients were negative (−0.0023 for 65–74 years and −0.0136 for 75–84 years) and were therefore set to zero. This approach has been used by others,[8] and is reasonable because several studies indicate non-significant differences in risk of coronary heart disease between smokers and nonsmokers over age 65.[12,21] Overall mortality and coronary heart disease incidence rates were based on US Vital Statistics.[22] Incidence rates for ages 35–74 were assumed to follow those observed in the Framingham Heart Study,[23] with a secular adjustment for the decline in coronary heart disease incidence since the beginning of the Framingham Heart Study,[24] and with extrapolation to ages 75–84 and linear interpolation to smooth the age-specific annual rates.[13]

Intervention

Using the interactive intervention feature of the model, we studied the effect of interventions that produce a substantial reduction in the number of male smokers in the year 1990 by 25%, 50%, and 75%, respectively, on the United States population in the year 2015, 25 years after the reduction was achieved. Interventions were assumed to apply equally across all age groups. The measures that were used to assess the impact of the smoking interventions were the percent change over expected baseline values (i.e., where no smoking intervention is modeled) that the smoking reductions would have on the following measures in the year 2015: a) size of population free of coronary heart disease, b) coronary heart disease incidence rates and c) absolute incidence (the total number of new cases) of coronary heart disease.

TABLE 1—Smoking Coefficients and Standard Errors Used in Demographic-Epidemiologic Model for Determining Risk of Developing Coronary Heart Disease

Age (years)	Model*	Framingham
35–44	0.0316 (0.015)	0.029 (0.010)
45–54	0.0203 (0.010)	0.015 (0.005)
55–64	0.0090 (0.008)	0.016 (0.004)
65–74	0.0 (0.011)	0.002 (0.006)
75–84	0.0 (0.017)	−0.040 (0.023)

*Model coefficients obtained from smoothing Framingham logistic regression function coefficients (see text for details).

Sensitivity Analyses

The variations in our results that would be caused by changes in the coefficients of the logistic risk function for developing coronary heart disease were determined by repeating our analysis four times, using different sets of coefficients each time. The first sensitivity analysis used the actual Framingham coefficients and allowed us to assess the impact that the smoothed coefficients had on our results. The second and third sensitivity analyses used the smoothed coefficients, but first increased and then decreased them by one standard error. These analyses allowed us to assess the effect that interventions on smoking would have if smoking were of more or less importance as a risk factor for coronary heart disease than originally assumed. The fourth analysis examined the impact that a 50 percent smoking reduction would have when smoking was considered as a risk factor for the development of coronary heart disease in the elderly. In this analysis, the smoothed coefficient for persons ages 55–65, 0.009 (Table 1) was applied to persons ≥65.

We also evaluated the impact that delaying the benefits of intervention programs by three years would have on our results. For these analyses, cessation programs begun in 1990 were assumed to have no effect until 1993.

Finally, we evaluated the impact of new smoking interventions on coronary heart disease incidence under the assumption that secular changes in smoking prevalence are already occurring independently of any new intervention programs.[25] For these analyses we estimated the incidence rate and absolute incidence of coronary heart disease in 2015 that would occur if current secular declines continued at a linear rate as they have since 1974[25] through 2000. These estimates were used to compute the incremental change in coronary heart disease incidence that would occur in 2015 if in 1990 a new intervention to reduce smoking by 25, 50, or 75 percent was combined with the projected secular decline through the year 2000.

Results

Increase in Population at-Risk

The projected long-term effect of each smoking intervention within each age group was to increase the number of men at risk for coronary heart disease (i.e. men alive who are free of coronary heart disease) (Table 2). These increases occur in large part because fewer men would die of smoking-related non-coronary diseases if there were reductions in smoking. The increases range from 54,565 to 301,925 persons, 25 years after a 50 percent reduction in smoking prevalence is achieved. Men ages 65–74 in 2015, who were ages 40–49 at the time of the intervention in 1990, had the largest percent increase.

Standardized Absolute Incidence

After adjusting for population differences, the overall reduction in the number of incident coronary heart disease cases in the year 2015, which is 25 years after the smoking reductions are achieved, is projected to range from 2.3% to 5.9% (Table 3). For ages 65–74, there are estimated to be small increases in the number of cases of coronary heart disease under some levels of smoking reduction, and for persons ≥75 years the number of cases is unchanged.

Incidence Rate

The projected baseline rate of coronary heart disease incidence in the United States male population in the year 2015, assuming that cohort-specific risk factor distributions

TABLE 2—Absolute and Percent Change over Baseline in Number of Males at Risk for Coronary Heart Disease (males who are alive and free of coronary heart disease) in the Year 2015 According to Level of Smoking Reduction in 1990

Age (years) in 2015	Level of Smoking Reduction					
	25%	(% Change)	50%	(% Change)	75%	(% Change)
35–44	+28,897	(+0.2)	+54,565	(+0.3)	+77,292	(+0.4)
45–54	+108,846	(+0.6)	+208,283	(+1.1)	+298,800	(+1.6)
55–64	+155,498	(+1.0)	+301,925	(+1.9)	+439,344	(+2.8)
65–74	+91,865	(+1.1)	+180,753	(+2.2)	+266,534	(+3.2)
75–84	+31,681	(+1.0)	+62,881	(+2.0)	+93,567	(+2.9)
Total	+416,787	(+0.7)	+808,407	(+1.3)	+1,175,537	(+1.9)

TABLE 3—Percent Change Over Baseline in Age-Standardized Absolute Incidence of Coronary Heart Disease among Males Free of Coronary Heart Disease According to Level of Smoking Reduction in 1990 (Numbers in Table 2 are obtained by applying incidence rates that occur under each intervention to the baseline population observed in 2015 when no smoking interventions are modeled.)

Age (years)	Level of Smoking Reduction		
	25% (% Change)	50%	75%
35–44	(−9.7)	(−16.7)	(−23.6)
45–54	(−5.2)	(−8.3)	(−13.0)
55–64	(−1.2)	(−2.6)	(−4.0)
65–74	(−0.1)	(+0.4)	(+0.4)
75–84	(0.0)	(0.0)	(0.0)
35–54	(−6.6)	(−10.8)	(−16.2)
35–84	(−2.3)	(−3.9)	(−5.9)

change with age as they have historically, resulted in coronary heart disease incidence rates between 0.29 and 1.96 per 100 (Table 4A). In the year 2015, 25 years after each level of smoking cessation is achieved, there were large decreases in the coronary heart disease incidence rates in the younger age groups. However, a 25 percent reduction in the number of smokers would have no influence on coronary heart disease incidence rates in males over age 65. Larger percent reductions in the number of male smokers would actually *increase* the coronary heart disease incidence rate among 65–74 year olds, while leaving coronary heart disease incidence among 75–84 year olds unchanged. The reasons for these increases are as follows: a) the pool of elderly survivors has been enriched with non-smokers who have slightly less favorable cholesterol levels, blood pressures, and weights; b) many of the at-risk smokers have already developed coronary heart disease; and c) Framingham data indicate that smoking is not a risk factor for coronary heart disease in the elderly (Table 1).

Absolute Incidence

The projected percent change over baseline estimates in the annual absolute incidence of new coronary heart disease cases (i.e. total number) 25 years after each level of smoking cessation is achieved is given in Table 4B. Smoking inter-

TABLE 4—Projected Percent Changes (in parentheses) over Baseline Values for CHD Incidence Rate, Absolute CHD Incidence, and Total CHD Prevalence, 25 Years after each level of Smoking Reduction is Achieved

A. Change in CHD Incidence Rate (per 100 among males free of CHD)

Age (years)	Baseline Incidence Rate	Change in Incidence Rate for Each Level of Smoking Reduction		
		25%	50%	75%
35–44	0.29	(−10)	(−17)	(−24)
45–54	0.64	(−5)	(−8)	(−13)
55–64	1.42	(−1)	(−3)	(−4)
65–74	1.81	(0)	(+1)	(+1)
75–84	1.96	(0)	(0)	(0)
35–54	0.47	(−6)	(−11)	(−16)
35–84	0.96	(−2)	(−4)	(−5)

B. Change in Absolute CHD Incidence (among males free of CHD)

Age (years)	Absolute Baseline Incidence	Change in Absolute Incidence for Each Level of Smoking Reduction		
		25%	50%	75%
35–44	50,471	(−9)	(−16)	(−22)
45–54	119,111	(−4)	(−8)	(−11)
55–64	219,786	(0)	(−1)	(−1)
65–74	149,072	(+1)	(+2)	(+4)
75–84	63,171	(+1)	(+2)	(+3)
35–54	169,582	(−5)	(−10)	(−15)
35–84	601,611	(−1)	(−2)	(−3)

*CHD: Coronary heart disease

ventions did not reduce absolute incidence of coronary heart disease as greatly as they reduced coronary heart disease incidence rates. In particular, the total number of new coronary heart disease cases *increases* for males over age 65. This is because absolute incidence reflects the size of the population at risk, which increases for all age groups as a result of smoking interventions.

Sensitivity Analyses

Changes in Smoking Coefficients—Changes in the smoking coefficients would alter estimates of coronary heart disease incidence rates and absolute incidence equally in younger men. However, under all of our sensitivity analyses (Table 5), the absolute incidence of coronary heart disease would increase in men over age 65. For example, if the actual raw, unadjusted Framingham logistic risk function coefficients were used rather than the smoothed coefficients, the results changed only to a minor degree. The largest difference was in men ages 75–84, in whom the Framingham coefficients are least stable. In this age group, the projected increase in the absolute incidence of coronary heart disease would be even more dramatic (4% higher) with the raw coefficients.

If smoking were not as strong a risk factor for coronary heart disease (i.e. if coefficients were decreased by one standard error), coronary heart disease incidence rates would still decrease for males under age 65, and the effects on absolute coronary heart disease incidence would be approximately one-half as large as under the baseline assumptions. In contrast, if smoking were a more powerful risk factor for coronary heart disease (i.e. if coefficients were increased by one standard error), a 50 percent reduction in the number of smokers in 1990 would result in reductions in coronary heart disease incidence rates by 2015 for *all* age groups. However, under both these scenarios the absolute incidence of coronary heart disease would increase in men over age 65.

When smoking was modeled as a risk factor for development of coronary heart disease in persons ≥65, incidence rates declined in the older age groups. However, absolute incidence still increased (by +1% rather than +2%) in persons ≥65.

Lag Time—Because smoking cessation programs may not produce immediate benefits, the analysis was repeated under the assumption that there is a three-year lag before benefits from smoking cessation would be realized. When the effects of smoking cessation programs begun in 1990 were delayed until 1993, the results changed by less than 1 percent.

Consideration of Future Secular Changes—When the analysis considered the current secular changes in smoking prevalence and assumed that rates of change would continue through the year 2000,[25] the incidence rates and absolute incidences in 2015 were different from those that would pertain if current secular trends do not continue. For example, a continuation of current secular rates of change would be about equivalent to a new intervention that yielded a 25 percent reduction in smoking in 1990.

When the secular decline through 2000 was used as the baseline for estimating the effects of new intervention programs to reduce smoking rates in 1990 by an additional 25, 50, or 75 percent, the effects of the new programs were dampened, but followed the same patterns previously reported. For example, the incremental impact of a 50 percent reduction in the number of smokers in 1990 when combined with a continuation in the secular decline through 2000 was less than the incremental impact of a 50 percent reduction in smoking rates if there are no secular declines, but greater than the incremental impact of a 25 percent reduction in smoking rates if there are no secular declines independent of

TABLE 5—Percent changes[*] over baseline values (i.e., no smoking intervention) for incidence rate and absolute incidence of CHD under alternate logistic risk function coefficients when a 50% decrease in smoking is achieved in 1990.

A. Change in Incidence Rate of CHD in 2015

	Coefficients Used in Logistic Risk Function				
Ages (years)	Smoothed Coefficients %	Unadjusted Coefficients %	Smoothed −1 S.E. %	Smoothed +1 S.E. %	Smoothed Elderly[†] %
35–44	−17	−17	−7	−28	−17
45–54	−8	−6	−5	−14	−8
55–64	−3	−5	−1	−5	−3
65–74	+1	0	0	−2	−2
75–84	0	+5	0	−1	−2

B. Change in Absolute Incidence of CHD in 2015

	Coefficients Used in Logistic Risk Function				
	Smoothed Coefficients %	Unadjusted Coefficients %	Smoothed −1 S.E. %	Smoothed +1 S.E. %	Smoothed Elderly[†] %
35–44	−16	−15	−8	−25	−16
45–54	−8	−5	−4	−12	−8
55–64	−1	−4	+1	−3	−1
65–74	+2	+2	+1	+2	+1
75–84	+2	+6	+1	+3	+1

[*] Note that changes are calculated relative to baseline simulations using the type of coefficients specified. For example, the −28% change in the incidence rate for 35–44 year olds that is observed when smoothed +1 S.E. coefficients are used, was calculated relative to baseline simulations that used the smoothed +1 S.E. coefficients. [†] Smoothed elderly refers to analyses done using smoothed coefficients (Table 1), with coefficients in persons ≥65 of 0.009 (the same coefficient as in the 55–64 age range).

new interventions for ages 35–64. For ages ≥65 years, the incremental impact of a 50 percent reduction when combined with a secular decline was identical to the impact of a 50 percent reduction when no secular declines were modeled.

Discussion

It is important to emphasize that because smoking is a risk factor for the development of many debilitating diseases, including chronic obstructive pulmonary disease and lung cancer, the ability of smoking cessation to improve the overall health of the population is clear. Our study is distinct from other studies, which have sought to measure the effects of risk factor interventions on coronary heart disease mortality and incidence,[3,5,6,21,26–28] because we have examined the long-term, steady-state, population wide impact that smoking cessation can have in the primary prevention of coronary heart disease among US males ages 35–84 who are free of coronary heart disease.

Our results indicate that interventions at a variety of levels would reduce the incidence rate of coronary heart disease for males under age 65, but that incidence rates would increase or remain unchanged for males ≥65 years of age. Note that although our baseline model, which was based on Framingham data, did not consider smoking to be a risk factor for the development of coronary heart disease among males ≥65 years,[12] males in this age group in 2015 were ages 40–59 at the time of the smoking intervention. Thus, they benefited from smoking reductions for at least six years before reaching age 65.

The fact that smoking interventions had the largest impact in reducing coronary heart disease incidence in the younger age groups is consistent with reports in the literature from several populations to indicate that smoking is a stronger relative risk factor for coronary heart disease in the young than in the elderly. The Framingham Heart Study reported incidence rates showing that non-smokers had a higher incidence of coronary heart disease than did smokers over the age of 65, however these differences were not statistically significant.[29] A study of British physicians by Doll and Peto[21] reported similar findings. Hennekens, *et al*,[30] noted that although the relative risks of coronary heart disease associated with smoking may decrease with age, the absolute risk increases. This phenomenon may be explained by the higher background risks of coronary heart disease that are associated with aging, and because many of the "at-risk" smokers may have already died from coronary heart disease or other smoking-related causes.[30,31]

A more recent analysis of data from the Framingham Heart Study indicated that smoking was of borderline significance as a risk factor for coronary heart disease in persons age 65 and older after adjusting for systolic blood pressure, left ventricular hypertrophy, total serum cholesterol, blood glucose, and Metropolitan relative weight.[31] A study of men under age 55 showed decreased risks of myocardial infarction after quitting smoking.[3] A similar phenomenon was reported for elderly men who quit smoking, but the endpoint was coronary heart disease mortality rather than myocardial infarction and this study included men with preexisting coronary heart disease.[5,6] On the basis of these reports and a prospective study that identified smoking as a risk factor for the development of new coronary heart disease in the elderly,[32] we undertook a sensitivity analysis, which modeled smoking as a risk factor in persons ≥65. Although decreases in incidence rates were observed for males ≥65, the intervention still impacted more substantially on the younger age groups.

One striking finding from our analysis is the small *increase* observed in the incidence of coronary heart disease among males ≥65 years of age that occurs under some of the smoking interventions, an effect that was observed even when smoking was modeled as a risk factor in the elderly. These increases emphasize that smoking interventions delay, but do not prevent, the development of coronary heart disease in the male population.

For all smoking interventions, decreases in absolute incidence were less pronounced than were decreases in the incidence rates for males <65 years, because of the larger population at-risk for coronary heart disease as a result of the decrease in smoking-related deaths under each intervention. In addition, there were increases in absolute incidence ranging from 1–4 percent for males ≥65 years under various scenarios. Thus, although decreases in the number of cases of new coronary heart disease would result from smoking interventions for younger persons, there would be a net increase in the number of incident cases of coronary heart disease for males ≥65 years of age.

On the basis of data from the literature, Goldman and Cook[33] estimated that the short-term reduction in coronary heart disease mortality due to an approximate 15 percent decrease in the number of smokers would be between 4.2 and 6 percent. Our results are somewhat different because they represent the differences observed in coronary heart disease incidence and absolute incidence in all men ages 35–84 free of heart disease 25 years after the smoking reduction was achieved. The overall standardized long-term impact of smoking interventions among males aged 35–84 was to reduce the total number of incident coronary heart disease cases by 2.3 percent when smoking rates were reduced 25 percent, by 3.9 percent when smoking rates were reduced by 50 percent, and by 5.9 percent when smoking rates were reduced by 75 percent. These reductions are less than might have been expected.

The Multiple Risk Factor Intervention Trial,[26] which studied men ages 35–57 over a six-year period, estimated that their observed 14 percent difference in smoking rates between the usual care and special intervention groups had the potential to produce a 5 percent difference in the number of coronary heart disease events. Our results for men ages 35–54 at 25 years after a 25 percent reduction in smoking rates was achieved indicate an age-standardized 6.6 percent difference in incident coronary heart disease cases. This estimate reflects the longer-term steady-state projected impact of smoking cessation on coronary heart disease incidence.

Our model depends in part on the accuracy of the National Vital Statistics data on which it is based. Although such data indicate a continued decline in ischemic heart disease (ICD Codes 410–414) mortality since 1980, a good portion of this decline is offset by increases in deaths due to diagnoses such as cardiac arrest and congestive heart failure.[22] These trends suggest that changes in coding practices since the institution of the revision of the ICD criteria in 1975[34] may have had an effect on the decline in deaths that are classified as due to ischemic heart disease.

In addition, the prevalence of ischemic heart disease, as estimated from the updates of the Health Interview Survey,[35] has increased by more than 10 percent since 1980. These trends, in the face of declining coronary risk factors, are consistent with more aggressive, early diagnosis of coronary disease rather than a true increase in atherogenesis. These issues indicate that the Coronary Heart Disease Policy Model will always be addressing a moving target, and thus we plan periodic updates as sufficient epidemiologic data on coronary disease incidence, prevalence, deaths and risk factor distribution become available, usually with a lag time of five or more years.

We did not explicitly consider the economic implications of smoking cessation. Oster, *et al*,[8] have predicted that the increased expenditures attributable to longer life-expectancy under smoking cessation programs are likely to be similar to the savings obtained from reductions in smoking-related illnesses, and Warner[7] has pointed out that the quality of life for non-smokers is likely to be higher than that of smokers. It is therefore not surprising that a variety of smoking interventions have been shown to be cost-effective.[11,36] Warner[7] also stated that although the incidence of cardiovascular disease would decrease in a smoke free society, the age-unadjusted incidence might increase. Our results provide further evidence in support of this conjecture. We conclude that although smoking interventions will not decrease the rate of coronary heart disease incidence for all age groups, the reduction in the number of cases of coronary heart disease among younger persons is substantial, and will offset the slight increases in the older age groups.

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Overcoming Potential Pitfalls in the Use of Medicare Data for Epidemiologic Research

ELLIOTT S. FISHER, MD, MPH, JOHN A. BARON, MD, MSC, DAVID J. MALENKA, MD,
JANE BARRETT, MSC, AND THOMAS A. BUBOLZ, PHD

Abstract: We used Medicare data bases and US Census data to address two questions critical to the use of Medicare files for epidemiologic research. First, we examined the degree to which the population enrolled in the Medicare program is similar to the elderly resident population of the United States, as estimated by the US Census. We found small differences in the total population estimates but substantial differences by age and race. Second, we found that among Medicare enrollees, physician claims identify a small proportion of hip fracture cases which are not documented in the hospital discharge files. This proportion varies by age, region, and state within the United States. Calculation of rates based on Medicare hospital discharge data, and probably other hospital discharge data sets as well, must take these limitations into account. Use of all available Medicare data files can overcome these limitations. (*Am J Public Health* 1990; 80:1487–1490.)

Introduction

Administrative health care data bases, such as those maintained by the Health Care Financing Administration (HCFA) for the Medicare program, provide an increasingly accessible and widely used source of data for health care research. Medicare files have been used to study disease incidence and risk factors,[1–3] to document regional variations in treatment patterns,[4,5] to study hospital-associated mortality rates and quality of care,[6] and to evaluate the outcomes of specific surgical and medical treatments.[7,8]

Most concern about the validity of research findings based on the Medicare files has focused on issues related to the accuracy of diagnostic coding.[9–11] In this paper we highlight several other potential limitations of the Medicare data relevant to epidemiologic research by addressing two questions. First, can one assume that the population captured in Medicare files is the same as the elderly population of the United States? Second, do the Medicare hospital claims data identify all discharge records for Medicare enrollees who were hospitalized for a specific disease? The answers to these questions highlight how the careful use of Medicare data can overcome these limitations.

Methods

Population Estimates

We used US Census Bureau survey data and Medicare data files to estimate the age, sex, and race-specific counts of United States residents and Medicare enrollees who were over the age of 65 and were alive on July 1, 1985. Census estimates were taken from the Current Population Estimates.[12] Estimates for the Medicare eligible population were determined by using the 5 percent National HISKEW (live) file maintained by the Health Care Financing Administration. This file is a random sample of the Medicare population, based on the terminal two digits of the social security number and records the birth date, state of residence, race, gender, and Medicare eligibility status of beneficiaries who have been eligible for Medicare at some point and are not known to have died. (Deaths are recorded in a different file.) We excluded Medicare eligibles listed as residing outside the 50 States. We then determined age, sex, and race-specific counts of individuals in the HISKEW files generated on 12/31/84 and on 12/31/85. We averaged these counts and multiplied by 20 to estimate the Medicare population alive on July 1, 1985.

Hip Fracture Case Ascertainment

We searched the 5 percent National MEDPAR (hospital discharge) files for all discharges in calendar year 1985 with a discharge diagnosis of hip fracture (ICD-9-CM code 820.0–820.9) and the 5 percent Medicare Part B (physician and other provider) files for all claims specifically documenting the treatment of an acute hip fracture during calendar year 1985 (Current Procedure Terminology codes in the range 27230 to 27248). We developed a detailed algorithm to define acute hip fractures based on all claims submitted during the calendar year for potential cases, relying on both diagnosis and procedure codes and the dates of the related claims.[13] To be included, cases had to have at least one hospital claim specifying the diagnosis, or two Part B claims supporting the diagnosis of hip fracture. We excluded potential cases for whom the only evidence of treatment for the fracture was provided by a single Part B claim (3 percent of candidate cases). This decision was based on extensive review of all claims for a sample of these potential cases. Hip fracture cases were excluded if there was evidence that the treatment was provided for an earlier fracture or that the fracture resulted from metastatic cancer. We applied this algorithm to the data obtained from the 1985 5 percent national sample of Part A and Part B claims.

Results

Population Estimates

We compare the US Census estimate of the elderly population with our estimate of the Medicare population in Table 1. Overall, the Medicare population, as estimated from the HISKEW file, represents approximately 96 percent of the US Census estimate of the elderly population, but the proportion enrolled in Medicare varies sharply by both age and race. Among Whites, only 90 percent of the estimated elderly population between the ages of 65 and 69 is enrolled in Medicare and the proportion enrolled increases with age. Among Blacks, only 79 percent of the estimated population

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From the Department of Medicine, Dartmouth Medical School (Fisher, Baron, Malenka); Department of Community and Family Medicine at Dartmouth (all authors); and the Department of Veterans Affairs Medical Center, White River Junction, Vermont (Fisher). Address reprint requests to Elliott S. Fisher, MD, MPH, Department of Community and Family Medicine, Dartmouth Medical School, Hanover, NH 03756. This paper, submitted to the Journal May 9, 1990, was revised and accepted for publication August 12, 1990.

TABLE 1—Comparison of Population Estimates based on Medicare Enrollment File (HISKEW) and US Current Population Survey (counts in thousands)

Race/Age	Male				Female				All			
	HSK	CNS	H/C	C-H	HSK	CNS	H/C	C-H	HSK	CNS	H/C	C-H
White												
65–69	3425	3821	0.90	396	4196	4615	0.91	419	7620	8436	0.90	816
70–74	2705	2899	0.93	194	3674	3914	0.94	240	6380	6813	0.94	433
75–79	1797	1924	0.93	127	2880	3044	0.95	164	4677	4968	0.94	291
80–84	1021	1051	0.97	30	1998	2001	1.00	3	3019	3052	0.99	33
80+	674	695	0.97	21	1760	1779	0.99	19	2434	2474	0.98	40
Total	9623	10390	0.93	767	14508	15313	0.95	805	24131	25743	0.94	1612
Black												
65–69	287	361	0.79	74	374	472	0.79	98	661	833	0.79	172
70–74	230	258	0.89	28	331	373	0.89	42	561	631	0.89	70
75–79	152	172	0.89	20	250	270	0.93	20	403	442	0.91	39
80–84	87	84	1.04	−3	164	149	1.10	−15	251	233	1.08	−18
85+	65	65	0.99	0	146	140	1.04	−6	210	205	1.03	5
Total	821	940	0.87	119	1265	1404	0.90	139	2086	2344	0.89	258
Other Race												
65–69	65	72	0.91	7	79	89	0.89	10	144	161	0.90	17
70–74	53	56	0.95	3	59	66	0.90	7	113	122	0.92	9
75–79	36	39	0.91	3	31	46	0.67	15	66	85	0.78	19
80–84	20	19	1.05	−1	18	26	0.70	8	38	45	0.84	7
85+	11	13	0.81	2	14	19	0.75	5	25	32	0.77	7
Total	185	199	0.93	14	201	246	0.82	45	386	445	0.87	59
Unknown/missing race												
65–69	118	0	—	−118	155	0	—	−155	273	0	—	−273
70–74	97	0	—	−97	130	0	—	−130	227	0	—	−227
75–79	62	0	—	−62	106	0	—	−106	168	0	—	−168
80–84	28	0	—	−28	55	0	—	−55	83	0	—	−83
85+	16	0	—	−16	57	0	—	−57	74	0	—	−74
Total	322	0	—	−322	503	0	—	−503	825	0	—	−825
All Races												
65–69	3895	4254	0.92	359	4803	5176	0.93	373	8698	9430	0.92	732
70–74	3086	3213	0.96	127	4194	4353	0.96	159	7281	7566	0.96	285
75–79	2047	2135	0.96	88	3268	3360	0.97	92	5315	5495	0.97	180
80–84	1156	1154	1.00	−2	2235	2176	1.03	−59	3391	3330	1.02	−61
85+	766	773	0.99	7	1978	1938	1.02	−40	2743	2711	1.01	−32
Total	10951	11529	0.95	578	16478	17003	0.97	485	27428	28532	0.96	1104

Notes: Estimates of Medicare enrolled population calculated by taking average of live counts for 5% national sample for 12/31/84 and 12/31/85 in each relevant cell and multiplying by 20. Persons recorded as on the live HISKEW file but who were not residents of the 50 states were excluded. Census estimates were for 7/1/85, and were obtained from the Current Population Survey.[12]

HSK = Estimate of population from Medicare's HISKEW file.
CNS = Estimate of population from Current Population Survey.
H/C = Ratio of HISKEW estimate to CPS estimate.
C-H = Census estimate minus HISKEW estimate

aged 65 to 70 is enrolled in Medicare. Among Blacks, the proportion enrolled increases with age and exceeds the Census estimates among those over 80 years old. No notable differences between the two genders in the age-specific proportions enrolled in Medicare are apparent for Whites or Blacks. However, among those whose race is designated as "other," different patterns are apparent for males and females. Finally, for 3.1 percent of the individuals in the HISKEW file, race is designated as unknown, while the Census assigns a race designation to all individuals.

Case Ascertainment from Hospital Claims

Table 2 presents data on the proportion of hip fracture cases identified from the Part B (physician and other provider) files by state for 1985. Overall, 9.1 percent of hip fracture cases were identified from the Part B claims alone, and the proportion of cases so identified ranged from none in several states to a high of 32.8 percent in Virginia. The proportion of cases identified only through the Part B claims decreased significantly with age from 10.6 percent for those ages 65 to 69 to 8.4 percent for those age 85 and over (Chi square, 4df, 12.9, p = .01). However, no significant differences were found in the proportion of cases identified through the Part B files by gender or race.

Discussion

It is often assumed that the Medicare-eligible population corresponds closely to the elderly population of the United States and that studies of Medicare's hospital discharges will provide accurate estimates of hospital utilization among the elderly. One recent study has estimated hip fracture incidence rates using data from the US Census to estimate age/race/gender-specific denominators, while using Medicare hospital discharge data to calculate age, race, and gender-specific numerators for these rates.[3] Our data suggest that this approach must be taken with considerable caution.

We found that the Medicare population over the age of 65 differs in several important ways from the US Census estimates of the elderly population as a whole. The fact that only 95 to 96 percent of US elderly residents are enrolled in Medicare has been previously reported,[14] but breakdowns by age, gender, or race were not provided. Our data demonstrate that Medicare coverage varies substantially across these groups. Although some of the differences between Census and Medicare estimates of the population are trivial, others

TABLE 2—Proportion of Hip Fracture Cases Identified through the Medicare Part B Claims only, by State, based on 5% National Sample, Calendar Year 1985

State/Region	N Cases	% B Only
NORTHEAST		
Connecticut	128	0.0
Maine	46	4.3
Massachusetts	249	13.3
New Hampshire	38	7.9
New Jersey	276	18.1
New York	685	11.1
Pennsylvania	479	7.9
Rhode Island	48	12.5
Vermont	19	0.0
Region Total	1968	10.6
MIDWEST		
Illinois	439	11.4
Indiana	262	7.6
Iowa	156	4.5
Kansas	132	3.0
Michigan	268	9.7
Minnesota	151	4.6
Missouri	254	4.7
Nebraska	63	4.8
North Dakota	33	9.1
Ohio	443	11.5
South Dakota	32	3.1
Wisconsin	208	2.4
Region Total	2441	7.7
SOUTH		
Alabama	141	10.6
Arkansas	109	5.5
Delaware	27	3.7
District of Columbia	9	11.1
Florida	522	11.7
Georgia	186	7.5
Kentucky	159	6.9
Louisiana	99	4.0
Maryland	111	9.9
Mississippi	78	7.7
North Carolina	179	27.4
Oklahoma	146	4.8
South Carolina	68	5.9
Tennessee	195	4.6
Texas	521	7.5
Virginia	131	32.8
West Virginia	78	6.4
Region Total	2759	10.4
WEST		
Alaska	3	0.0
Arizona	107	5.6
California	821	8.6
Colorado	107	9.3
Hawaii	10	0.0
Idaho	33	3.0
Montana	29	3.4
Nevada	24	0.0
New Mexico	43	2.3
Oregon	123	5.7
Utah	27	0.0
Washington	163	6.7
Wyoming	15	6.7
Region Total	1505	7.2

are not. Consideration of sampling variations would further complicate interpretation of these differences. However, our analyses underline the obvious but important point that Medicare data provide information about the Medicare population; findings of studies that rely on Medicare data should be generalized to the elderly population as a whole with caution.

A full exploration of the causes of these differences is beyond the scope of this paper but would include issues related to criteria for Social Security eligibility, delayed enrollment for those still employed after age 65, and decisions by eligible persons about whether to enroll in either or both of the Medicare Part A and Part B programs. That the Current Population Estimates data from the US Census bureau are estimates subject to both sampling and measurement error may also play a role.

Equally important, coding conventions for race used by the Census and by Medicare differ. While the Census assigns a racial designation to all residents, approximately 3 percent of Medicare enrollees remain of unknown race. Given the White predominance in the elderly population and the possibility that substantial numbers of Whites are included in the group designated unknown, it would be inappropriate to group those of unknown race with either of the other two minority groups.

Finally, we must note that we cannot explain the relative excess of Medicare eligibles compared to Census estimates for the older age groups in several racial categories. Although Census undercounts in the oldest population are conceivable, this could also be explained by the failure to record a small number of deaths in the HISKEW file, a deficiency that would be magnified in the oldest age categories. We have found some indication that this might be the case. The 1985 5 percent national HISKEW file included 368 individuals over the age of 110, resulting in an estimated 7,360 individuals over this age in the population as a whole—an implausible figure. Nevertheless, given the close similarity of our overall estimates of the Medicare population based on HISKEW to HCFA's estimate of the proportion of the US population enrolled in Medicare (95 percent[14]), few deaths overall are apparently missed by these files and analytic difficulties caused by this problem would be serious only among the oldest old.

While it is obvious that outpatient claims would be required to analyze incidence rates for fractures that can be treated in an outpatient setting, we found that Medicare's hospital discharge files do not capture all treatment for hip fractures, a condition virtually always treated on an inpatient basis. Overall, physician claims were required to identify 9.1 percent of hip fracture cases; almost all of these physician claims were coded with the place of service "inpatient." For the majority of these cases, we found associated hospital claims with evidence of treatment for orthopedic conditions related to the hip, but no diagnosis of hip fracture. For the remainder, we must conclude that the hospital claims were either lost in the data transmission process or severely miscoded. It is reassuring that only small and insubstantial differences in fracture detection rates were found by age, race or gender, but the variation by state of residence might pose problems for analyses of geographic differences in incidence rates across states. We should also note that gaps in the physician claims are possible—Connecticut Part B claims were missing from the 1985 Medicare Part B 5 percent national file. A second caveat is that some Medicare enrollees may receive care in Veterans' Administration hospitals. We found that approximately 4 percent of New England hip fracture cases among males (and less than 1 percent among females) were documented in VA files.[13]

Our data illustrate two specific limitations beyond the accuracy of diagnostic coding that must be considered when using or reporting findings of studies that rely on Medicare data. First, the Medicare population cannot be assumed to correspond exactly with the resident US elderly population. Second, gaps in the Medicare hospital data must be acknowledged. These potential limitations could also apply to any analysis where different data sources are used for the numerator and denominator, or where hospital discharge data sets are not complete prior to release as public use tapes.

Most important, however, our data underline the feasibility of overcoming these limitations through the careful linkage of data from the multiple files available through the Medicare program. One of the cardinal rules in calculating rates is to ensure that the criteria used to determine eligibility for inclusion in the numerator should be as similar as possible to those used for assignment to the denominator. The Medicare enrollment file (HISKEW) can be used not only to define precisely denominators by age, race, sex, and mailing address (zip code), but can also be used to define identical inclusion criteria for both the numerator and denominator of rates. Finally, physician claims obtained from the Part B files and hospital claims from the Veterans' Administration can be merged with the Medicare hospital claims files to reduce the likelihood of missing cases in the study population of interest. When the limitations of administrative health care data bases are recognized and appropriate care exercised in their use, they can provide a valuable tool for epidemiologic research.

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Prospective Payment and the Utilization of Physical Therapy Service in the Hospitalized Elderly

PETER HOLT, MD, AND CAROL HUTNER WINOGRAD, MD

Abstract: We studied the effect, in a university teaching hospital, of the prospective payment system (PPS) on utilization of physical therapy (PT), a non-reimbursable service; subjects were hospitalized patients aged 75 or older with non-PT-related diagnoses (myocardial infarction, pneumonia, congestive heart failure, and colectomy) and PT-related diagnoses (cerebrovascular accident and hip fracture). The proportion of patients referred for PT increased from 68 percent pre-PPS to 85 percent post-PPS for those with PT-related diagnoses and from 13 percent pre-PPS to 19 percent post-PPS for those with non-PT-related diagnoses. The mean number of sessions of PT decreased slightly for both groups: from 8.5 to 7.6 sessions for those with PT-related diagnoses and from 5.2 to 4.5 for those with non-PT-related diagnoses. In patients with PT-related diagnoses whose ambulatory status worsened during hospitalization, referrals for PT increased from 76 percent pre-PPS to 98 percent post-PPS. Referrals of comparable patients with non-PT-related diagnoses did not increase. Changes in provider education and efforts to reduce length of stay may account for these findings. (*Am J Public Health* 1990; 80:1491–1494.)

Introduction

In 1983, the Health Care Financing Administration (HCFA) introduced the prospective payment system (PPS). PPS created an incentive for hospitals to provide care at a cost at or below the established amount for a specific diagnosis-related group (DRG). Cost reduction could be accomplished by either reducing length of stay (LOS) or by decreasing utilization of ancillary services, such as physical therapy (PT). Nationally, LOS has decreased since the implementation of PPS.[1–5] In contrast, little attention has been directed toward the impact of PPS on utilization of PT.

This study focused on PT because it is a service used by hospitalized patients with a broad range of diagnoses. In 1986, approximately one in six hospitalized patients received some type of PT treatment.* Second, PT is a clinical service designed to improve mobility, ambulation, and independence. Because ambulatory status affects both quality of life and discharge status, the accessibility to and availability of PT is important, particularly to the frail elderly.[6–8]

Previous studies examining the effects of PPS on utilization of hospital PT services have shown mixed results. Several studies revealed a significant decrease in the mean number of PT sessions in hip fracture patients treated after the implementation of PPS.[2,5,9] These studies, however, did not evaluate the percentage of patients referred for PT. Another study found a significant increase in referrals for inpatient PT since 1983 in small acute care rural hospitals.[10] In contrast, a more recent study from 501 non-federal acute care hospitals reported a decrease in the number of patients referred for PT since 1983.[11] Neither of the latter two studies addressed which categories of patients (i.e. which DRGs) were more likely to have experienced changes in the pattern of utilization of PT.

The current study was designed to document the utilization patterns of PT services in the elderly prior to and after the implementation of PPS via two different indicators of PT utilization: percentage of referrals and mean number of sessions.

Methods

Medical records of patients 75 years of age or older discharged from Stanford University Medical Center (SUMC) were examined retrospectively for six consecutive years from September 1, 1980 through August 31, 1986. This represented four years prior to the implementation of PPS (9/1/80–8/31/84) and two years after the implementation of PPS (9/1/84–8/31/86).

Subjects

All patients discharged with one of the following six principal conditions were eligible: cerebrovascular accident (CVA); pneumonia; myocardial infarction (MI); congestive heart failure (CHF); hip fracture; and colectomy. These were further categorized into two groups: PT-related diagnoses which included CVA and hip fracture, and non-PT-related diagnoses which included pneumonia, MI, CHF, and colectomy. MI was not included in PT-related diagnoses because Medicare does not recognize it as a "rehabilitation diagnosis" in spite of the proliferation of cardiac rehabilitation programs.

After PPS was implemented, there was an incentive for chart abstractors in the Medical Records Department to be as specific as possible when coding. In the two years post-PPS, fewer patients than pre-PPS were discharged with CVA. Instead they were categorized with the specific etiology of CVA (e.g. embolic, thrombotic, or hemorrhagic) which resulted in a different ICD code. All other categories remained the same pre- and post-PPS.

The study sample (n = 917) was selected from a list of patients by use of tables of random permutations.[12] Because of the small numbers in the CVA category for the post-PPS years, all eligible patients were included.

Data Collection and Analysis

Data were abstracted from the medical records with the number of PT sessions counted only if the patient actually received therapy. The initial PT evaluation was counted as one session. The percentage of referrals and the mean number of sessions were calculated.

From St. Luke's Hospital, Kansas City Missouri (Holt); Department of Medicine, Division of Gerontology, Stanford University School of Medicine, California (Winograd); and the Veterans Affairs Medical Center, Palo Alto, California (Holt, Winograd). Address reprint requests to Peter Holt, MD, St. Luke's Hospital, 44th and Wornall Road, Kansas City, MO 64111. This paper, submitted to the Journal May 15, 1989, was revised and accepted for publication May 29, 1990.

*Terbush TW: Estimates of physical therapy costs and usage by Medicare patients in FY 1986. Prepared for the American Physical Therapy Association, January 1987.

Ambulatory status was categorized as follows: unable to walk, walks only with the physical assistance of another person, ambulates independently with or without ambulation aids, or not obtainable. It was determined just prior to the admission illness or injury (i.e. pre-morbid ambulatory status) and at discharge. For those patients in whom ambulatory status was obtainable at both times, the difference between pre-morbid ambulatory status and discharge ambulatory status was scored better, worse, or the same.

Residence at admission and discharge was categorized as follows: independent home setting, board and care home, acute care facility, intermediate care facility (ICF), skilled nursing facility (SNF), or other type of facility. Any change in type of residence was recorded.

The predictor variable was relationship to PPS, divided into pre-PPS and post-PPS. The outcome variables were utilization of PT (measured by mean number of PT sessions and percent of patients referred to PT), length of stay, change in ambulatory status, and change in type of residence.

Results were analyzed using t-tests for the continuous outcome variables and chi-square tests for dichotomous outcome variables. Ninety-five percent confidence intervals were calculated for all results.[13]

Results

We identified 917 patient charts for review after random selection. Of these, 907 (99%) were available for review, 468 (98%) of the 475 cases in the pre-PPS and 439 (99%) of the 442 cases in the post-PPS period. The two groups were similar in terms of age, sex, marital status, type of residence, and in-hospital mortality (Table 1). Length of stay decreased by 1.6 days.

The proportion of patients referred for PT increased from 68 percent pre-PPS to 85 percent post-PPS (diff = 17%, CI = 7.4, 26.6%) for the PT-related diagnoses group and from 13 percent pre-PPS to 19 percent post-PPS (diff = 6%, CI = 0.3, 11.7%) for the non-PT-related diagnoses group. However, the mean number of PT sessions decreased slightly from 8.5 sessions pre-PPS to 7.6 sessions post-PPS (diff = 0.9, CI = −0.8, 2.6%) for the PT-related diagnoses group and from 5.2 sessions pre-PPS to 4.5 sessions post-PPS (diff = 0.7, CI = −0.7, 1.4%) sessions post-PPS for the non-PT-related diagnoses group (Table 2).

Patients with PT-related diagnoses whose ambulatory status worsened were referred for PT more often post-PPS (76% pre-PPS to 98% post-PPS, diff = 22%, CI = 9.5, 34.5%), while PT referrals of comparable patients with non-PT-related diagnoses showed no change. Discharge to a more dependent level of care increased for patients with PT-related diagnoses (44% pre-PPS to 59% post-PPS, diff = 15%, CI = 2, 28%), while no difference existed in the group with non-PT-related diagnoses.

TABLE 1—Demographic Data of Subjects Pre- and Post-Prospective Payment System

Item	Pre-PPS (N = 468)	Post-PPS (N = 439)	Difference (95% CI)
Age (Mean)	82.5	83.0	0.5 (−0.2, +1.2)
Sex (% Male)	38.0	38.0	0
Marital Status (% Married)	38.0	38.0	0
Residence (% Home)	86.0	84.0	2 (−13, +17)
Mortality (%)	16.0	13.0	3 (−11, +17)
Length of Stay (Number of Days)	10.7	9.1	1.6 (+0.7, +2.5)

The proportion of patients ambulating independently prior to admission did not change for the group with PT-related diagnoses however, there was a small decrease at discharge (54% pre-PPS to 46% post-PPS, diff = −8%, CI = −5, 21%). The proportion of patients in the non-PT-related diagnoses group that ambulated independently decreased prior to admission (95% pre-PPS to 90% post-PPS, diff = −5%, CI = 0.2, 9.8%) and at discharge (87% pre-PPS to 80% post-PPS, diff = −7%, CI = 0.03, 13.7%) (Table 3).

Discussion

This study documents an increase in the percentage of patients referred for PT after the implementation of PPS in a university hospital. This is a surprising and encouraging finding, because prior to the study we expected to find a decrease in utilization of PT services. Our findings do not substantiate previous concerns that the elderly would inevitably suffer from poorer quality of care under PPS, as hospitals under pressure to maintain their "bottom line" would reduce the provision of potentially beneficial services.

A small decrease occurred in the mean number of PT sessions administered in the post-PPS time period. This decrease was less than one might have expected given the considerable decrease in the LOS; this contrasts with the results of other investigators who found decreases in the LOS and mean number of PT sessions to be more closely related.[2,5,9]

Physicians, who must initiate the referral for PT, may have become more aware of the benefits of early mobilization and ambulation during acute illness; referral for PT services may help maintain mobility and thereby enable patients to be discharged sooner.

The overall ambulatory status of patients appears to have declined during the period of implementation of PPS. After PPS, patients with non-PT-related diagnoses (e.g. CHF and colectomy) were more dependent in ambulation both prior to admission and at discharge. These patients were admitted with medical and surgical conditions that in themselves did not directly affect neurological or musculoskeletal function and should not have had an impact on ambulation. This dependency in ambulation may reflect the findings of other investigators that the case mix of patients after PPS changed.[14] Acute care hospitals now provide care to more severely ill patients in need of the greatest intensity of service. For example, after PPS, patients with mild forms of pneumonia and CHF may have been treated as outpatients, while only the most debilitated patients were admitted to the hospital.

Since one of the primary indications for and goals of PT is to maintain ambulation, we also examined the change in ambulatory status of patients admitted before and after PPS, and the referral patterns for PT in these patients. Although no difference occurred in the percentage of patients whose ambulatory status worsened during the hospitalization before or after PPS, subgroup analysis revealed striking differences between PT-related and non-PT-related groups in referral for PT. Patients with strokes and hip fracture whose ambulation worsened were more frequently referred for PT after PPS. In contrast, comparable patients with medical and surgical diagnoses experienced no increase in referrals for PT after PPS.

TABLE 2—Utilization of Physical Therapy Services Pre- and Post-Prospective Payment System by Diagnostic Group

	PT-Related Diagnoses			Non-PT-Related Diagnoses		
	Pre-PPS	Post-PPS	Difference (95% CI)	Pre-PPS	Post-PPS	Difference (95% CI)
Percent Referred	68	85	+17 (+7.4, +26.6)	13	19	+6 (+0.3, 11.7)
Mean Number of PT Sessions	8.5	7.6	−0.9 (−0.8, +2.6)	5.2	4.5	−0.7 (−0.7, +1.4)

TABLE 3—Percent of Patients Ambulating Independently Pre- and Post-Prospective Payment System by Diagnostic Group

	PT-Related Diagnoses			Non-PT-Related Diagnoses		
	Pre-PPS	Post-PPS	Difference (95% CI)	Pre-PPS	Post-PPS	Difference (95% CI)
Pre-Morbid (Prior to Illness Causing Hospitalization)	90	90	0	95	90	−5 (+0.2, +9.8)
At Discharge	54	46	−8 (−5, +21)	87	80	−7 (+.03, +13.7)
Pre-Post Difference	−36	−44	+8 (−5, +21)	−8	−10	+2 (+3, +7)

Impaired mobility is a common problem in the elderly[15] that is frequently exacerbated by acute care hospitalization.** Simple bedrest, a common order on admission for acutely ill patients, can be devastating in the frail elderly.[16–18]

We reviewed a random selection of 12 patient charts from the non-PT-related group with a decline in ambulatory status, but were not referred for PT services. No consistent factor (e.g. patient refusal, severity of illness, physician preference) could be identified to explain the absence of a PT request. Referrals may be based on different standards of practice among groups of physicians. Neurologists and orthopedists may be more familiar with rehabilitation than are internists. One would hope that patients with impaired mobility regardless of their diagnosis would receive the necessary PT services prior to discharge.

An encouraging finding in our study was that the percentage of patients with non-PT-related diagnoses requiring placement in a more dependent level of care did not increase after PPS. This is corroborated by other investigators.[19,20] However, more patients in the PT-related group (e.g. hip fractures) required a more dependent placement after PPS. Several studies have reported similar findings;[2,5,9] change in case mix, as well as a decreased LOS may explain them. After PPS, patients with hip fractures and strokes may not have had adequate recovery time in the hospital to ambulate well enough to go back to their original type of residence.

We do not think the decrease in the number of patients with CVA diagnostic code, after the implementation of PPS, affected the comparability of the two groups because the change in coding specified only the type of stroke and not the location or associated functional disability. Both groups most likely had an equal need for utilization of PT services.

In 1984, the Palo Alto Veterans Administration Medical Center, a major teaching hospital for Stanford University Medical Center housestaff, established a geriatrics program with a strong rehabilitation focus for acutely ill, hospitalized elderly veterans. Many members of the housestaff rotated on this service to learn the principles of geriatric medicine. These concepts may have been integrated into their treatment plans when the housestaff returned to Stanford Medical Center for subsequent rotations. This may have affected PT referrals.

A home health care program was introduced at Stanford during the study period. Greater availability of in-house support services, including health aides, nurses, and rehabilitation therapists, may have had an impact on discharge planning and discharge location.

In summary, this study describes a changing pattern of utilization of PT since the implementation of PPS in a university hospital. Specifically, the percent of patients referred for PT increased. In spite of this finding, this study also suggests that patients with medical and surgical diagnoses who have impaired ambulation may not be referred for potentially beneficial PT services.

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Conference Proceedings Available on Primary Care Research

As part of its emphasis on primary care for patients of all ages and conditions, the Agency for Health Care Policy and Research (AHCPR), in cooperation with several other agencies, convened a conference entitled "Primary Care Research: An Agenda for the 90s" in March 1990 at Colorado Springs, CO.

The conference explored many of the emerging issues expected to set the pace for primary care research during the 1990s, including: changing provider behavior, research with disadvantaged populations, medical treatment effectiveness in primary care mental health, community-oriented primary care, prevention research, outcome measurement issues, rural health, AIDS, substance abuse, practice-based research networks, and the federal perspective on primary care research.

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Public Health Briefs

Prevalence of Chronic Bronchitis among US Hispanics from the Hispanic Health and Nutrition Examination Survey, 1982–84

KI MOON BANG, PHD, MPH, PETER J. GERGEN, MD, MPH, AND MARGARET CARROLL, MSPH

Abstract: In the Hispanic Health and Nutrition Examination Survey (HHANES), Puerto Ricans had a higher age-adjusted prevalence of self-reported chronic bronchitis (2.9 percent, 95% CI = 2.2, 3.6) than Mexican Americans (1.7 percent, 95% CI = 1.3, 2.1) or Cubans (1.7 percent, 95% CI = 0.9, 2.5). The prevalence of chronic bronchitis was at least 2 times higher in smokers as compared to nonsmokers among Puerto Ricans and Cubans, but not for Mexican Americans. (*Am J Public Health* 1990; 80:1495–1497.)

Introduction

Chronic bronchitis is part of chronic obstructive pulmonary disease (COPD), the fifth leading cause of mortality in the US. Mortality from COPD increased 43 percent between 1979 to 1986 (22.2/100,000 to 31.8/100,000).[1] The epidemiology of chronic bronchitis has been described in numerous studies of US Whites.[2–4] Little is known about the prevalence of chronic bronchitis among the US Hispanic population.

We present estimates of the prevalence of self-reported chronic bronchitis among the three Hispanic subgroups and the association between chronic bronchitis and measured risk factors among Hispanics based on the 1982-84 Hispanic Health and Nutrition Examination Survey (HHANES).

Methods

The HHANES was conducted by the National Center for Health Statistics from 1982 to 1984. The target population consisted of individuals ages 6 months–74 years in the three major Hispanic subgroups in selected areas of the United States: Mexican Americans in five Southwestern states (Arizona, California, Colorado, New Mexico, and Texas); Cubans in Dade County, Florida; and Puerto Ricans in the New York City area. The sample design of HHANES was a multistage, stratified, clustered sample of each of the Hispanic subgroups. Details of the HHANES have been published.[5]

In the household interview, respondents were asked two questions about chronic bronchitis: 1) "Has a doctor ever told you that you had chronic bronchitis?", 2) "Do you still have chronic bronchitis?" Positive responses to both questions were used to define chronic bronchitis.

All statistical analyses were carried out using programs compatible with the Statistical Analysis System.[6] Since HHANES was based on a complex survey design, sampling weights were used to estimate the prevalence of chronic bronchitis and complex samples variances were estimated by multiplying the simple random sample variance estimates by an average design effect.[7,8] To compare the prevalence of chronic bronchitis among the three Hispanic subgroups, estimates of chronic bronchitis were age-standardized to the distribution of the 1980 population. The trend for the age-specific prevalence was analyzed using the Armitage method.[9] The logistic regression analyses were performed using RTILOGIT program.[10] Odds ratios (ORs) and 95% confidence intervals (CI) of ORs are presented.[11]

Results

The response rates were 79.2 percent for Mexican Americans, 73.9 percent for Cuban Americans, and 80.3 percent for Puerto Ricans (Table 1). The level of nonresponse did not vary across age, race, sex, or income groups. The age distributions of the analytical sample were similar in the Mexican American and Puerto Rican samples, but the Cuban American sample was considerably older. The proportion of heavy smokers (> 20 cigarettes/day) among Puerto Rican and Cuban smokers was much higher than among Mexican American smokers (Table 1).

The age-adjusted prevalence of self-reported chronic bronchitis was 1.7 percent (95% CI: 1.3, 2.1) for Mexican Americans; 1.7 percent (95% CI: 0.9, 2.5) for Cuban Americans; and 2.9 percent (95% CI: 2.2, 3.6) for Puerto Ricans (Table 2); prevalence increased slightly with age in both males and females among Mexican Americans ($p < 0.001$). Failure to find a clear increase in prevalence with age among Cuban Americans and Puerto Ricans may be due to the smaller sample size of these groups.

To assess the role of measured risk factors of self-reported chronic bronchitis, logistic models were fitted using age, sex, smoking, and poverty status (Table 3). Former smokers were excluded from the analysis because of small numbers. In the model for Puerto Ricans, age (OR = 1.03, 95% CI = 1.01, 1.05), sex (OR = 2.69, 95% CI = 1.38, 5.26), and cigarette smoking (OR = 2.72, 95% CI = 1.57, 4.71) were significant risk factors for chronic bronchitis. However, only smoking (OR = 2.75, 95% CI = 1.14, 6.62) for Cuban Americans and age (OR = 1.03, 95% CI = 1.01, 1.05) for Mexican Americans were significantly associated with chronic bronchitis.

From the National Center for Health Statistics, Hyattsville, MD. Address reprint requests to Ki Moon Bang, PhD, MPH, Division of Health Examination Statistics, Room 900, National Center for Health Statistics, 6525 Belcrest Road, Hyattsville, MD 20782. This paper, submitted to the Journal February 20, 1990, was revised and accepted for publication May 29, 1990.

TABLE 1—Description of the Samples

Survey period	Mexican Americans July 1982–November 1983	Cuban Americans January–April 1984	Puerto Ricans May–December 198
Number of sample persons selected	6,868	1,874	2,748
Response rate (%)	79.2	73.9	80.3
Age Groups (%)			
12–34 yrs	60.1	37.9	56.8
35–54 yrs	26.1	34.8	28.0
55–74 yrs	13.8	27.3	15.2
Sex			
(%) Male	46.7	46.2	41.5
(%) Female	53.3	53.8	58.5
Cigarette smoking (%)			
Current smokers	27.2	26.3	29.8
Former smokers	14.8	17.4	12.6
Never smokers	58.0	56.3	57.6
% of heavy smokers (>20 cig./day)	23.8	58.3	50.0
Poverty income ratio (%)			
Below poverty (<1)	31.3	20.7	48.0
At/Above poverty (≥1)	68.7	79.3	52.0

TABLE 2—Age-specific Prevalence (%) of Chronic Bronchitis among Hispanic subgroups

	Males/Age (years)			Females/Age (years)			
	12–34	35–54	55–74	12–34	35–54	55–74	Total **
Mexican Americans							
Number	1,578	624	336	1,688	798	416	5,440
Prevalence	0.9	1.5	3.6	0.9	2.5	3.1	1.7
95% CI	0.7, 1.1	0.4, 2.7	1.8, 5.4	0, 1.8	1.4, 3.7	0.8, 5.4	1.3, 2.1
Cuban Americans							
Number	256	221	162	268	260	217	1,384
Prevalence	*	3.5	2.3	1.5	2.7	1.8	1.7
95% CI	*	0.3, 6.7	0, 5.3	0, 3.3	0.2, 5.2	0, 4.8	0.9, 2.5
Puerto Ricans							
Number	552	229	135	701	389	201	2,207
Prevalence	1.1	2.0	2.5	2.7	5.0	4.9	2.9
95% CI	0.2, 2.0	0, 4.1	0, 7.3	1.3, 4.1	2.5, 7.5	1.9, 7.9	2.2, 3.6

CI: Confidence interval.
*No reported chronic bronchitis.
**Age-adjusted prevalence.

Discussion

The high age-adjusted prevalence of chronic bronchitis among Puerto Ricans may have been due to the higher proportion of heavy cigarette smokers among Puerto Ricans and/or the high exposure to air pollution in New York City. Females had consistently higher age-adjusted prevalences of chronic bronchitis than males: in Puerto Ricans, 3.8 percent for males, 1.6 percent for females; in Mexican Americans, 1.8 percent for females, 1.6 percent for males; in Cuban Americans, 1.9 percent for females, 1.4 percent for males. A recent study on 2,111 Hispanics in New Mexico reported that the prevalence of chronic bronchitis among females was 4.3 percent as compared to 2.9 percent for males.[12] In the 1978 National Health Interview Survey, the prevalence of chronic bronchitis was higher in females (3.7 percent) than in males (2.9 percent).[13] A possible explanation for the female predominance might be that females visit clinics more often and are therefore aware of this diagnosis.[14] In contrast, some studies reported that the risk of chronic bronchitis was more than three-fold greater for men than women.[15]

The failure to find an association between smoking and chronic bronchitis in Mexican Americans may be due to the high proportion of light smokers. Other studies have shown cigarette smoking to be one of the most important risk factors for chronic bronchitis.[16,17] A recent study[18] found that heavy smokers had a six-fold increase in chronic bronchitis as

TABLE 3—Estimates of Odds Ratios with 95% Confidence Intervals of Risk Factors for Chronic Bronchitis[a]

	Odds Ratio	95% Confidence Interval
Mexican Americans		
Sex	1.09	0.66, 1.82
Age	1.03	1.01, 1.05
Smoking	1.04	0.61, 1.77
Poverty	0.84	0.50, 1.39
Cuban Americans		
Sex	1.28	0.53, 3.10
Age	1.02	1.00, 1.04
Smoking	2.75	1.14, 6.62
Poverty	0.61	0.24, 1.57
Puerto Ricans		
Sex	2.69	1.38, 5.26
Age	1.03	1.01, 1.05
Smoking	2.72	1.57, 4.71
Poverty	0.57	0.33, 0.99

Sex (1 = female, 0 = male), Age (continuous)
Smoking (1 = smokers, 0 = nonsmokers), Poverty (1 = below, 0 = at/above)
a) Past smokers excluded.

compared to nonsmokers. Other important risk factors of chronic bronchitis, not measured in HHANES, include occupational exposures,[19-21] airborne irritant gases,[22,23] viral or bacterial infection,[4,24] and familial aggregation.[25]

The American Thoracic Society's definition of chronic bronchitis requires a history of three months of cough and phlegm or sputum for more than two consecutive years.[26] This information was not collected during HHANES. The use of diagnosed cases only will underestimate the prevalence of chronic bronchitis. Other potential biases which may occur cannot be estimated in the data set but must be kept in mind when interpreting the results of this study.

In conclusion, our results show that chronic bronchitis was more prevalent in Puerto Ricans than Cuban Americans or Mexican Americans. Cigarette smoking was an important risk factor for chronic bronchitis among Puerto Ricans and Cuban Americans. Thus, smoking prevention and smoking cessation programs in the Hispanic population would help reduce the risk of chronic bronchitis.

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Prevalence and Restrictiveness of Smoking Policies in King County, Washington, Manufacturing Worksites

SUSAN KINNE, PHD

Abstract: This study examined prevalence of smoking policies and of significant smoking restrictions in a 1988 random sample of 1,528 manufacturing companies in the Seattle, Washington area, and related these characteristics to worksite size and industry type. Sixty-nine percent of companies had formal smoking policies and 85 percent regulated smoking to some degree. Workforce size was directly related to prevalence of significant smoking restrictions, but when type of industry was controlled, the linear trend in the size/prevalence relationship was sustained in only three of the 18 industry groupings. (*Am J Public Health* 1990; 80:1498–1500.)

Introduction

Policies regulating workplace smoking have been imposed to reduce fire hazard, protect equipment[1] and, most recently, to protect nonsmokers from exposure to environmental tobacco smoke.[2] Efforts to assess the effects of these policies on smoking reduction and cessation have been reported.[3–8] Little is known about policy prevalence and content, because most policy surveys have had nonrandom sampling techniques, undefined populations, and low response rates that limit the quality of the data.[9–11] Few studies have clearly distinguished among types of policies or investigated relationships between organizational characteristics and policy content.[12] This research was designed to provide an accurate assessment of the prevalence of degrees of policy restriction on smoking in a geographically defined population of organizations employing many smokers, and to describe the relationship between policy, organizational size, and industry type.

Methods

These data are from a random sample of all King County (Washington) manufacturing worksites, surveyed by telephone between November 1988 and July 1989. King County includes the city of Seattle. The sampling frame, the set of all companies with manufacturing Standard Industrial Codes (SIC),[13] and King County zip code mailing addresses, was identified by a commercial broker (n = 2,883). Two-thirds of these companies (n = 1,922) were randomly selected for the study. Three hundred ninety-four companies were deleted as ineligible (disconnected numbers; non-County or non-manufacturing concerns; and one-person companies) on grounds that such companies would not use policy to control behavior. The response rate among eligible companies was 91.3 percent (1,395/1,528).

Companies in the sample received a letter describing the study, followed by a 10-minute telephone interview with the company head or a person identified by the head as knowledgeable about smoking policy. The informant was asked, "Does your organization presently have a policy on where employees can and cannot smoke in your buildings and facilities?" Policy was defined as a directive on smoking by the company head or governing body, communicated to employees in written or unwritten form.

Policies vary in the degree to which they limit smoking. Table 1 gives the policy response categories in order of increasing restrictiveness. Worksites also can be grouped by their degree of policy-mandated protection against environmental tobacco smoke (ETS) and the inferred company attitude toward smoking. Companies with no policy, with non-restrictive policies, or with smoking allowed except in no-smoking areas offer little ETS protection and accept smoking as the norm.[14] These were labeled "Smoking Allowed." Companies in which nonsmoking is the norm[14] were grouped as "Significant Restrictions." These include worksites that limit smoking to designated smoking areas and those that ban smoking completely, thereby limiting ETS exposure substantially. The prevalence of "Significant Restrictions" was examined for worksites of differing workforce size and industry type.

TABLE 1—Distribution of Smoking Policy Restrictions by Number of Employees in King County (WA) Manufacturing Worksites

Policy Restrictiveness	Number of employees 2–10	11–25	26–100	>100	Total %/n
(n of companies in category)	(700)	(326)	(269)	(100)	(1395)
Smoking allowed					
No policy on smoking	19.7%	15.6	4.5	1.0	14.5/202
Policy allows smoking everywhere	0.0	0.6	0.4	1.0	0.3/4
Non-policy smoking restrictions	19.6	14.7	11.9	10.0	16.3/227
Smoking except in no-smoking areas	4.9	9.2	14.1	10.0	8.0/112
(subtotal: smoking allowed)	(44.1)	(40.2)	(30.9)	(22.0)	(39.1/545)
Significant restrictions					
Smoking only in designated areas	12.9	26.7	40.1	57.0	24.5/342
No smoking allowed	27.9	27.0	23.4	16.0	25.9/362
No smoking, prefer to hire nonsmokers	10.9	4.9	4.5	4.0	7.7/108
No smoking, only hire nonsmokers	4.3	1.2	1.1	1.0	2.7/38
(subtotal: significant restrictions)	(55.9)	(59.8)	(69.1)	(78.0)	(60.9/850)

Address reprint requests to Susan Kinne, PhD, Fred Hutchinson Cancer Research Center, MP-702, 1124 Columbia Street, Seattle, WA 98104. This paper, submitted to the Journal January 29, 1990, was revised and accepted for publication May 29, 1990.

TABLE 2—Relationship between Prevalence of Significant Smoking Restrictions and Workforce Size, Controlling for SIC Group, King County (WA) Manufacturing Worksites

Industry by SIC Group Number (ranked by mean N of employees per company)	N of companies in group	Prevalence of restrictions (%) by N of employees				Prevalence for Group (%)	χ^2 Test for Trend	(p)
		2–10	11–25	26–100	> 100			
37: Transportation equipment	85	41.7	46.2	57.9	76.5	52.9	5.396	.020
36: Electrical, electronic equiment	84	63.6	53.8	77.8	90.9	70.2	3.425	.064
20: Food and kindred products	75	58.8	87.5	76.2	76.9	76.0	.711	.399
22: Textile mill products	9	60.0	—	66.7	100.0	66.7	.333	.564
29: Petroleum refining	9	33.3	100.0	100.0	100.0	77.8	3.493	.062
26: Paper and allied products	32	60.0	77.8	85.7	83.3	75.0	1.538	.215
38: Measuring, photographic, clocks	46	65.4	77.8	88.9	100.0	73.9	2.661	.103
33: Primary metal industries	28	27.3	75.0	28.6	0.0	39.3	.015	.901
30: Rubber and plastic products	50	50.0	62.5	70.0	100.0	64.0	4.132	.042
23: Apparel and fabric products	50	77.3	53.8	46.2	100.0	64.0	2.053	.152
34: Fabricated metal products	129	34.5	27.3	64.5	28.6	39.5	3.090	.079
35: Machinery, except electrical	138	44.2	57.1	47.8	90.0	50.7	3.815	.051
39: Miscellaneous manufacturing	101	55.9	54.2	100.0	85.7	62.4	4.679	.031
24: Lumber and wood products	67	62.9	70.6	50.0	100.0	64.2	0.025	.875
25: Furniture and fixtures	44	57.1	53.8	75.0	100.0	61.4	1.002	.317
32: Stone, clay, glass, concrete	48	36.4	47.1	71.4	0.0	43.8	.827	.363
28: Chemicals and allied products	41	59.1	55.6	90.0	—	65.8	1.948	.163
27: Printing, publishing	355	65.2	65.8	76.0	75.0	67.0	1.492	.222
31: Leather goods	4		—	—	—	75.0	—	—
Total companies	1,395					60.9%		

Results

We found that 69.2 percent (966/1,395) of companies reported having a policy on where employees could smoke in buildings and facilities (Table 1). Four of these companies (0.3 percent) stated that they had adopted a smoking policy that explicitly permitted smoking in all parts of the worksite, leaving 68.9 percent (962/1,395) of companies with policies that restricted smoking. An additional 227 companies without smoking policies reported that smoking was not allowed in parts of their worksite. When the latter two categories were combined, the data show that smoking was limited to some degree in 85.2 percent (1,189/1,395) of responding worksites. Companies that banned smoking completely, with or without hiring policies on smokers, comprised 36.3 percent (508/1,395) of worksites surveyed. The prevalence of significant restrictions on smoking was 60.9 percent (850/1,395).

The number of full-time employees, a measure of company size, was related to smoking policy (Table 1). Having a formal smoking policy increased with workforce size, as did prevalence of significant smoking restrictions, although the likelihood of a total smoking ban was greatest in the smallest companies. Companies without smoking restrictions were in the minority in every size category.

This relationship between workforce size and prevalence of significant restrictions varied by type of manufacturing activity. When SIC grouping was controlled, the chi-squared test for trend for the relationship between size and prevalence was significant for only three of the 18 SIC groups (Table 2). Five groups showed weak size/prevalence relationships ($.05 < p < .10$), not all of which can be attributed to small numbers of cases. In the remaining 10 groups, the size/prevalence relationships were erratic, suggesting that factors peculiar to those industries may account in part for the distribution of significant smoking restrictions.

Discussion

The results of this survey demonstrate general acceptance of the concept of worksite smoking control among King County manufacturers of all sizes. The majority of employees in these industries now enjoy some protection from exposure to environmental tobacco smoke.

These data share some of the limitations of previous work on smoking policy. We have no measures of reliability or validity of the questions used to inquire about policy, and respondents' meaning of "a policy on smoking" may vary.[12] Although the survey data estimate the extent of policies, they provide no information on policy enforcement, making it impossible to estimate actual impact on air quality. Lastly, these data reflect practices of manufacturers in one geographic area, and may not be representative of this sector's smoking policies in other parts of the country. A number of prominent employers and local governmental units have "gone smoke-free" in King County in the past few years. Washington State, King County, and the City of Seattle impose relatively stringent restrictions on public smoking. None of these jurisdictions requires that private employers or worksite environments have policies (although such a law, exempting small worksites, has been proposed by the State Legislature), but public acceptance of smoking restriction is widespread. This attitude may explain the survey's finding that extensive control of smoking in the worksite may be accomplished without legislative mandate.

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Depressive Symptoms and Cigarette Smoking among Latinos in San Francisco

Eliseo J. Pérez-Stable, MD, Gerardo Marín, PhD, Barbara V. Marín, PhD, and Mitchell H. Katz, MD

Abstract: We administered the Center for Epidemiological Studies Depression (CES-D) Scale to 547 San Francisco Latinos as part of a random digit dialing telephone survey to evaluate smoking behavior. Both men and women current smokers had the highest mean CES-D levels (9.7 and 14.3, respectively). Logistic regression analysis adjusting for gender, acculturation, education, age, and employment showed that current smokers had an odds ratio of 1.7 (95% CI = 1.3, 2.2) for significant depressive symptoms compared to former smokers (OR = 1.1; 95% CI = .8, 1.6) and never smokers (OR = 1). (*Am J Public Health* 1990; 80:1500–1502.)

Introduction

Cigarette smoking continues to be the most common cause of preventable morbidity and mortality in the US.[1] Three community studies[2–4] and a study of adolescents[5] have reported on possible relations between smoking status and depressive symptoms with conflicting results.

Five population-based surveys of Latino communities in California have reported a high rate of depressive symptoms compared to non-Latino White samples.[6–11] Female gender,[6,8,9] younger age,[6,9] formerly married status,[9–11] lower income,[6,10,11] less education,[9,10] unemployment,[11] and inability to speak English[9] have been associated with a higher level of depressive symptoms. Factors associated with increased depressive symptoms in the Hispanic Health and Nutrition Examination Survey were female gender, low educational achievement, low income, and US birth combined with Anglo-oriented acculturation.[12] To date, no study has examined smoking status as it relates to depressive symptoms among Latinos.

Although the overall prevalence of cigarette smoking among US Latinos is somewhat lower than that for Whites or Blacks, this rate reflects the substantially lower smoking proportions of Latino women compared to men.[13–15] Increasing acculturation to the mainstream US culture results in more smoking by Latino women, but less among Latino men, and thus may eventually narrow this smoking gender gap.[15]

We examined the possible relation of depressive symptoms, cigarette smoking status, and demographic variables, within a community survey of Latinos in San Francisco, part of an ongoing evaluation of a community intervention to promote smoking cessation.

Methods

Subjects were sampled from the 27 census tracts in San Francisco with at least 10 percent Latinos and representing two-thirds (N = 55,541) of all Latinos (N = 83,373) living in San Francisco according to the 1980 US Census.[16] Telephone prefixes corresponding to the census tracts were identified using a reverse telephone directory and the Mitofsky-Waksberg[17] method for random digit dialing was applied.[18] This method identifies nonworking and other ineligible prefixes in a first stage.

A household was eligible if, on the initial question, the person answering the telephone self-identified as Latino or if the majority of residents were Latino. Within a given household, the Latino adult between 18 and 65 years of age who had most recently celebrated a birthday was invited to respond. A structured questionnaire was administered anonymously by trained bilingual interviewers after subjects gave verbal consent. Participants responded in Spanish (69.7 percent) or in English (30.3 percent); 5.5 percent of eligible participants refused to be interviewed.

Depressive symptoms were measured using the Center for Epidemiological Studies Depression scale (CES-D)[19] which has been applied in Spanish in other studies with Latinos.[6–8,11] A CES-D score of 16 or more is considered to represent significant depressive symptoms, and validation studies with psychiatric and community samples showed that approximately one-third of persons scoring 16 or more on the CES-D are diagnosed with clinical depression.[20,21]

Never smokers were defined as persons who had smoked less than 100 cigarettes in their entire life. Gender, age,

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Address reprint requests to Eliseo J. Pérez-Stable, MD, Division of General Internal Medicine, Department of Medicine, University of California, 400 Parnassus Avenue, Room A-405, San Francisco, CA 94143-0320. Dr. B. Marín and Dr. Katz are with that same department; Dr. G. Marín is with the Department of Psychology, University of San Francisco. This paper, submitted to the Journal August 18, 1989, was revised and accepted for publication April 9, 1990.

education, employment in last two weeks, birthplace, and national background were asked, but marital status was not determined. A five-item acculturation scale with excellent reliability and validity was administered and scores were dichotomized into low (1 to <3) and high (≥ 3 to 5) acculturation groups.[22]

Data were analyzed using a standard statistical software package for a microcomputer (BMDP).[23] Variables that have been shown in previous research to be associated with depression were entered into a logistic regression model along with smoking status: acculturation (low/high), age (<30 / ≥ 30 years), employment status (yes/no), and education (<12 years / ≥ 12 years). The categorical variable smoking status was coded into two variables: current smoking (yes/no) and former smoking (yes/no).

Results

A total of 604 Latinos agreed to be interviewed by telephone between June and August 1988; 57 (9.4 percent) were excluded from this analysis because of incomplete information; their demographic characteristics and smoking status were similar to those completing the CES-D and acculturation scales for gender, age, acculturation, birthplace, education, and smoking status, but they were more likely to be unemployed. The 547 subjects completing the CES-D had a mean age of 35 years (SD = 12), 11 years (SD = 4.1) of education, and 59 percent were women. The mean duration of time living in the US for the foreign-born was 12.4 years (SD = 9.5) and the mean acculturation score for the entire sample was 2.4 (SD = 1.2). Overall smoking prevalence age-adjusted by the direct method[24] to the 1980 US Census[16] was 19 percent, with more men (29 percent) smoking than women (13 percent).

The overall mean level of depressive symptoms was 9.4 (95% CI = 8.6, 10.2), and 104 (19 percent) subjects scored 16 or greater on the CES-D. Table 1 shows the mean CES-D scores separately for men and women by demographic characteristics. Women reported a higher mean level of depressive symptoms than men for each category analyzed. Unemployment and less than high school education were also associated with a greater risk of depressive symptoms. Never smokers reported a lower mean level of depressive symptoms than current smokers or former smokers. The proportion of current smokers with a mean CES-D score ≥ 16 was 21.9 percent for men and 39.5 percent for women, compared to 11.8 percent and 18.5 percent, respectively, for never smokers.

The odds ratio (OR) of significant depressive symptoms (CES-D ≥ 16) by smoking status was evaluated using logistic regression analysis controlling for gender, acculturation, education, age, and employment status. Results, shown in Table 2, indicate that overall current smokers have a 70 percent greater risk of significant depressive symptoms than never smokers. The increased risk was present for both men and women. Former smokers do not have an increased overall risk of significant depressive symptoms although for women the risk is greater than for men.

Discussion

Among Latinos in San Francisco, current smokers have a modest increased risk of reporting significant depressive symptoms on the CES-D when compared to never smokers. Although unemployment and less years of education were associated with more depressive symptoms in this study, their effects were marginal compared to that of gender and smoking status. Marital status, social support and self-reported health status, which have been found to correlate with depressive symptoms in some studies,[9–11] were not determined as part of this survey and it is possible that these variables may partly account for our results.

TABLE 1—Mean Depressive Symptoms by Demographic Variables among San Francisco Latinos

	Men			Women		
	N*	CES-D	% $\geq$ 16	N	CES-D	% $\geq$ 16
Total	225	8.0	14.2	322	10.4	22.4
Age (years)						
18–29	96	8.1	13.5	117	10.2	22.2
30–65	129	7.9	14.7	204	10.4	22.5
Acculturation score						
low	146	8.4	15.1	226	10.5	22.6
high	79	7.2	12.7	96	10.1	21.9
Birthplace						
Latin America	172	8.0	14.5	234	10.3	21.4
United States	52	8.2	13.5	87	10.5	25.3
National background						
Mexican American	86	8.2	16.3	92	11.1	25.0
Central American	109	8.0	12.8	183	9.6	20.0
Other Latinos	30	7.3	13.3	41	11.9	26.8
Employed last 2 weeks?						
yes	165	7.4	12.1	183	10.0	20.2
no	57	9.8	21.1	138	10.9	25.4
Education (years)						
0–11	73	9.3	19.2	136	11.9	26.5
$\geq$12	150	7.3	12.0	183	9.3	19.7
Smoking status						
never	110	6.9	11.8	242	9.3	18.5
former	51	8.2	9.8	37	12.9	27.0
current	64	9.7	21.9	43	14.3	39.5

*Ns may not add up to total because of missing data.

Studies of personality profiles of smokers have generally found that smokers are more extroverted and that a modest

TABLE 2—Adjusted[*] Odds Ratio of Significant Depressive Symptoms by Cigarette Smoking Status among San Francisco Latinos[†]

	Sample OR (95% CI%)	Men OR (95% CI)	Women OR (95% CI)
Never smokers[‡]	1.0	1.0	1.0
Former smokers	1.1 (.8,1.6)	.8 (.5,1.5)	1.3 (.9,2.0)
Current smokers	1.7 (1.3,2.2)	1.4 (.9,2.2)	1.8 (1.2,2.5)

[*] Estimates were adjusted for acculturation level, education, age, and employment status. Estimates for the entire sample were also adjusted for gender.

[†] Significant depressive symptoms are defined as a CES-D score ≥16 (see text).

[‡] Never smokers were the reference group.

relationship exists between smoking and neuroticism.[25] The few specific studies on smoking and depression have found contradictory results. In a survey of 1,003 adults living in Los Angeles in 1979, Frerichs and colleagues[2] also found a significantly higher level of depressive symptoms as measured by the CES-D among smokers compared to nonsmokers (10.02 vs 8.76). In a study of 2,018 British men evaluated with a scale for psychoneurotic profile, Haines, *et al*,[3] also found that scores on the depression subscale (and the other four) were higher for smokers than nonsmokers. However, in both the above mentioned studies, when confounding factors were controlled the association of smoking and depression was no longer significant. A community survey of 1,209 men and women in New Zealand, using a scale similar to the British study, did find that depression was significantly greater among heavy smokers when compared to never smokers even when controlling for confounding factors.[4] Finally, a sample of 50 patients undergoing elective cardiac catheterization in St. Louis were evaluated for clinical depression and cigarette smoking was the only predictor variable associated with a diagnosis of depression.[26]

Although the causality between feeling depressed and smoking cannot be assumed from these cross-sectional data, recent prospective studies support such a link. Depressed adolescents 15 to 16 years of age in New York public schools were found to be more likely to be heavy cigarette smokers than non-depressed adolescents at ages 24 to 25 years.[5] Other suggestive evidence is from a clinical trial of clonidine as a pharmacologic adjunct to smoking cessation in 71 heavy smokers, which showed that 61 percent had a history of major depression and that this history had a significant negative effect on cessation regardless of treatment.[27]

It is possible that smokers report more depressive symptoms than nonsmokers because they are more willing to admit negative characteristics about themselves. As smoking becomes a less socially acceptable behavior smokers may become more depressed as a consequence, thus confounding the association observed in this study. However, if these findings are confirmed in other studies, smoking cessation interventions may need to integrate depression prevention and treatment components to maximize effectiveness.

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Dental Office Practices for Tobacco Users: Independent Practice and HMO Clinics

HERBERT H. SEVERSON, PHD, ELIZABETH G. EAKIN, BS, VICTOR J. STEVENS, PHD, AND EDWARD LICHTENSTEIN, PHD

Abstract: Surveys of independent dental practitioners and health maintenance organization (HMO) dentists and their adult male patients in Oregon found that cessation information and advice was being given to tobacco-using patients, especially smokeless tobacco users, and that the tobacco users surveyed expressed interest in obtaining help to quit. Oral health providers expressed an interest in further training in how to help their tobacco-using patients to quit. (*Am J Public Health* 1990; 80:1503–1505)

Introduction

In 1985 it was estimated that over 12 million Americans ages 12 and older had used chewing tobacco or snuff (smokeless tobacco) in the past year[1]; among males who use smokeless, 43 percent used it almost daily.[2] Studies have shown correlations between smokeless tobacco use and gingivitis, gingival recession, leukoplakia, and oral cancers.[3]

In 1978, the House of Delegates of the American Dental Association (ADA) resolved that ADA members be encouraged "to undertake an educational effort to inform their patients of the systemic and oral health hazards of chewing tobacco and snuff." A 1964 ADA resolution had encouraged its members to advise their patients about the hazards of cigarette smoking.

Recent articles have reviewed dentists' involvement in smoking cessation counseling.[4,5] However, we are not aware of similar information about dentists' involvement with smokeless tobacco prevention. In this paper we report the results of two surveys of dentists, dental hygienists, and their adult male patients.

Methods

Independent Practice (IP) Survey

An explanatory letter was mailed to all practicing members of the Lane County Oregon Dental Society. Of the 114 dentists, 24 (21 percent) refused to participate and two were used as pilot offices. The study was conducted over a two-week period in 1988 during which time questionnaires were to be given to all male patients over 18 years of age visiting the dental office. Questionnaires were also completed by dentists and hygienists. The final questionnaire return rate for dentists was 83 percent (73/88). Data were also received from 65 hygienists and 806 adult male patients. Neither the base population of hygienists nor the population of eligible male patients was known.

HMO Survey

This study was conducted in the Kaiser Permanente Dental Care Program based in Portland, Oregon, which provided comprehensive dental care to 85,000 enrolled members. Seventy full-time dentists and 65 full-time dental hygienists were employed in seven fully staffed dental offices. The questionnaire return rate was 61 percent for dentists and 68 percent for hygienists.

Overall, 1,506 of age-eligible male patients (56 percent of the male patient population) completed the questionnaire. Observers making spot checks in the clinics noted that receptionists often failed to give the questionnaire to patients, but of those who did receive questionnaires more than 90 percent returned them to the reception desk as requested.

Results

Table 1 presents tobacco use and attitudes toward provider advice from each sample.

Table 2 summarizes the reported interactions with smokeless tobacco users and smokers for dentists in both settings. Most offer some kind of information and cessation advice. All types of information and advice were given more often with smokeless tobacco-using patients than with smoking patients. However, only a minority of providers have cessation materials available in their offices. Dentists tend to be more optimistic and comfortable when advising about smokeless use than about cigarette smoking. The majority of dental professionals in both surveys were interested in learning about a smokeless tobacco cessation program. The data from hygienists generally paralleled those of dentists.

The majority of dentists report being pessimistic about the effectiveness of their advice or the likelihood of a patient attempting a cessation clinic program (Table 3). The majority of hygienists reported the need for further training and patient resistance to referral as the most important obstacles.

Discussion

The results show that dentists and hygienists in both settings are concerned about patient tobacco use and inter-

TABLE 1—Male Patient Tobacco Use and Attitude toward Provider Advice

	Independent Practice (IP) N = 80%	HMO Clinics N = 1506
% Using smokeless	6	3.9
% Smoking	18	26.9
% Smoking and smokeless	1.4	0.7
Smokeless users interested in receiving cessation assistance	41	39
Smokeless users who would consider advice	27	33
Smokeless users reporting cessation advice	44	not asked

Address reprint requests to Herbert H. Severson, PhD, Research Scientist, Oregon Research Institute, University of Oregon, 1899 Willamette Street, Suite 2, Eugene, OR 97401. Ms. Eakin is Research Assistant at the Institute; Dr. Stevens is Investigator at Kaiser Permanente Center for Health Research; Dr. Lichtenstein is Research Scientist and Professor at the Institute and the University of Oregon. This paper, submitted to the Journal March 30, 1990, was revised and accepted for publication May 15, 1990.

TABLE 2—Independent Practice and HMO Dentists Reported Interactions with Smokeless Tobacco Users and Smokers

	IP Dentists (N = 73)		HMO Dentists (N = 43)	
	ST Users	Smokers	ST Users	Smokers
Advice and information usually given				
Discuss health hazards (%)	88	55	77	47
Discuss benefits of quitting (%)	62	51	42	35
Advised cut down use (%)	26	18	23	21
Advised quit cold turkey (%)	30	11	30	7
Prescribe nicotine gum (%)	8	5	2	5
No advice (%)	4	16	12	35
Perceived mean effectiveness of advice (7 pt scale)*	4	3	4	2
Feel comfortable in giving advice (7 pt scale)*	6	4	5	4
Have available information on quitting/health effects in office (%)	27	26	12	17

*1 = low, 7 = high

TABLE 3—Major Obstacles to Treating Tobacco Use*

	Independent Practice		HMO	
	Dentist	Hygienist	Dentist	Hygienist
Obstacles	%	%	%	%
I am pessimistic about people's ability to change	64	46	70	36
Patients resist referral to cessation clinics	56	52	56	61
I believe or patients believe intervention is not appropriate since tobacco use is a matter of personal choice	49	42	49	36
I have too little time	38	31	42	34
I need further training	37	55	33	52
There is lack of coordination between dentistry and tobacco cessation services	30	31	42	39
I have little confidence in effectiveness of available services	26	11	26	16
There are too few services to which patients can be referred	16	22	7	11

*Respondents asked to choose four of the eight statements listed.

ested in providing cessation information and/or advice. Unfortunately, dental professionals receive little training and have few materials available that would facilitate such actions. Dentists appear more comfortable giving advice to smokeless tobacco users than to smokers. The relevance of chew and snuff use to oral health is the most parsimonious explanation for this finding. The difference between smokeless tobacco use rates in the IP and HMO patient populations may be due to demographic differences, sampling differences, or personnel differences.

The patients surveyed reported receptiveness to cessation advice and, in fact, expect it from a dental professional. Direct feedback from a dentist or hygienist on oral health status, such as visually pointing to tobacco-associated oral lesions, gingival recession, or tooth staining, can produce a "teachable moment" wherein patients are receptive to cessation advice. This receptivity to advice from dentists and hygienists is supported by data from an in-depth interview study of smokeless tobacco users.[6]

Studies of physicians' advice to smoking patients have demonstrated modest but significant efficacy.[7–9] Recent reports indicate that dentists' direct advice also can have a significant effect.[11–14] Similar effects seem likely for smokeless tobacco. The obstacles to implementation include the need for specific cessation training for dentists and hygienists and self-help cessation materials for patients. Specific materials are now becoming available that will assist oral health professionals in getting more directly involved in advising patients who seek their help in order to quit the use of tobacco products.[14]

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Evidence for Gonococcal Transmission within a Correctional System

KAREN H. VAN HOEVEN, MD, WALLACE C. ROONEY, JR., MD, AND STEPHEN C. JOSEPH, MD, MPH

Abstract: In a study to examine sexually transmissable disease occurring within a large correctional system where sexual activity is prohibited, 27 male inmates acquired culture-proven gonorrhea from in-jail sexual activity during a three-month period. These results provide evidence to encourage inmate education about the acquired immunodeficiency syndrome (AIDS) and to support condom distribution programs in correctional facilities. (*Am J Public Health* 1990: 80:1505–1506.)

Introduction

Studies reveal that up to 20 percent of male inmates engage in homosexual activity at some time during their imprisonment.[1] Research protocols which explore homosexual activity in jails and prisons generally use a subjective, questionnaire-type approach, asking prisoners whether they have participated in homosexual activity before or during incarceration.[1–7] Objective evidence for sexual activity occurring during imprisonment has been presented only in a few case reports which describe outbreaks of syphilis or gonorrhea occurring in prisoners incarcerated for prolonged periods.[8,9]

Sexual activity places inmates at risk for acquiring sexually transmitted diseases—including gonorrhea and the acquired immunodeficiency syndrome (AIDS). The risk of infection with the human immunodeficiency virus (HIV) among New York City prisoners during homosexual activity with other inmates is even higher than that of the general population, since studies have shown these inmates to have higher rates of HIV infection at admission to prison than the general population.[10]

Sexual activity is prohibited in many correctional systems, including that of New York City. Thus, the provision of condoms to inmates often is deemed unnecessary and objectionable by correction officials, since condom distribution would imply that rules prohibiting sexual activity are not enforced, or are unenforceable.[11] Yet from a public health perspective, condom distribution to inmates is important to enable inmates to protect themselves from sexually transmitted diseases, including AIDS. This study was designed to answer the following question: is there **objective** evidence that sexual activity occurs **regularly** in a large jail system?

Methods

In 1986, the correctional facilities of the New York City Department of Correction had a capacity to house over 12,500 inmates, and admitted more than 73,000 prisoners. The average daily census was almost always 100 percent or more of capacity. Most of the prisoners were detainees, staying for only a short period of time while awaiting trial or bail. Each prisoner received an admission history and physical exmination which included collecting information concerning a history of gonorrhea, dysuria, or urethral discharge and performing a urinalysis. Any newly admitted inmate who became suspect for gonorrheal infection after admission screening was examined, cultured, and treated for gonorrhea if clinically indicated.

For the months of October, November, and December 1986, the names, prisoner identification numbers, and dates of positive culture for all male New York City inmates with culture-proven gonorrhea were collected. From the files of the New York City Department of Correction, the date of admission to the jail was obtained for these prisoners. If an admission date preceded the date of positive culture by more than one week, the inmate's chart was reviewed and pertinent data were collected. These data included: date symptomatic, collection date and source (urethral, anal, and/or throat) of the specimen, and history of previous gonorrheal infections and treatment received, if any.

Address reprint requests to Karen H. van Hoeven, MD, Department of Pathology, Albert Einstein College of Medicine, 1300 Morris Park Avenue, Bronx, NY 10461. At the time of the study she was with the Bureau of Medical Affairs, New York City Department of Health. Dr. Rooney is Medical Director, Prison Health Service, New York City Department of Health: Dr. Joseph was Commissioner, New York City Department of Health from 1986 through 1989. This paper, submitted to the Journal October 23, 1989, was revised and accepted for publication May 18, 1990.

HIV Seroprevalence among Male IVDUs in Houston, Texas

MARK L. WILLIAMS, PHD

Abstract: HIV seroprevalence was 8.4 percent in a sample of 921 heterosexual male intravenous drug users in Houston, Texas who were not in drug treatment at the time of the study. Males who were Black, injected drugs daily, or had a history of syphilis had greater odds of being HIV positive than participants without those characteristics. (*Am J Public Health* 1990; 80:1507–1509.)

Introduction

The association between acquired immunodeficiency syndrome (AIDS) and intravenous drug use is well documented.[1] The primary means by which intravenous drug users (IVDUs) become infected with the human immunodeficiency virus (HIV), the cause of AIDS, are known.[2,3] However, the risk of an IVDU becoming infected with HIV is not uniform across racial/ethnic groups nor geographic locations.[4–6] As well, rates of HIV seroprevalence have been reported to vary among IVDUs due to the drug injected, frequency of injection, use of shooting galleries, injecting with strangers, the number of sexual partners, and a history of sexually transmitted diseases.[7–11]

The purpose of this study was to examine HIV seroprevalence rates among intravenous drug using males recruited from their natural environments in Houston, Texas. The majority of studies reporting HIV seroprevalence rates among IVDUs have been conducted in drug treatment, detoxification, or some other clinical setting.[2] Few studies have reported on the seroprevalence rates among intravenous drug injectors not in treatment.

Methodology

Subjects for this study were recruited from Houston's intravenous drug using population not in drug treatment. Participants were asked to respond to an interviewer administered survey covering demographic characteristic, drug use, and sexual behaviors. IV drug users were recruited from the city's three predominant racial/ethnic groups by trained, indigenous outreach workers. Three neighborhoods with known high incidences of intravenous drug use were targeted for outreach activities using targeted sampling.[12,13] Although targeted sampling is purposive, it is not random.

Study participants were required to be 18 years, have injected an illicit drug at least once during the six months prior to participation, and not have been in drug treatment for 30 days prior to participation. Outreach workers successfully recruited 1,279 men willing to be interviewed. Of that number, 81 percent agreed to have blood drawn for HIV testing. Only those respondents with complete surveys and HIV testing were considered for inclusion in the data set. Twenty-three participants were excluded because of sexual relationships with men during the previous six months. The resulting study sample of 921 males was 65 percent Black, 24 percent White, and 12 percent Hispanic (primarily Mexican American).

All variables used in the analysis, except HIV serostatus, are self-reported. A respondent's racial/ethnic classification was determined by his stated racial/ethnic identity. Drug use variables analyzed were frequency of injection, length of injected drug use, injection in a shooting gallery, sharing needles with a stranger, number of persons with whom needles were shared, and drug(s) injected. Frequency of injection was categorized as less than weekly, weekly, or daily depending on how often a respondent reported injecting during the six months prior to the interview. Drugs injected were categorized according to the reported drug(s) a respondent said he had injected during the six months before the interview. Respondents who said they injected more than one drug could either have injected the drugs alone or as mixtures.

Variables related to sexual behaviors which were examined were a self-reported history of syphilis, gonorrhea, herpes, and chlamydia. The number of sexual partners a respondent reported during the six months prior to participating in the study was examined.

HIV serostatus was determined by ELISA and Western Blot assays. Blood samples which were repeatedly positive on the ELISA test were screened using Western blot analysis. Samples were classified HIV positive if protein bands P24, P31, and EP41 or EP160 were present.

Univariate relationships were estimated using odds ratios and 95% confidence intervals. Significance was confirmed using Fisher's exact test (not shown).[14] To test the independence of significant univariate associations, a model of those relationships was analyzed using maximum likelihood method of logistic regression.[14,15] Coefficients are reported in independent odds ratios.

Results

HIV infection rates varied by race/ethnicity. As shown in Table 1, Black males had a seroprevalence rate of 10 percent. Seven percent of the White males were HIV positive. Black males were 2.80 times as likely to be HIV positive as Hispanic males.

Frequency of injection was found to be related to seropositivity. Less than weekly users had a 4 percent seropositive rate (Table 1). Eight percent of weekly users were HIV positive. Daily users had the highest number of positive test results, 11 percent. Associations between seroprevalence and self-reported history of syphilis, gonorrhea, herpes, and chlamydia were examined. Only a self-reported history of syphilis was significantly related to seroprevalence. Males who reported having had syphilis were 2.34 times as likely to be HIV positive as those who had not.

To test the independent association between race/ethnicity, frequency of injection, a past history of syphilis, and HIV seroprevalence logistic regression analysis was used. As shown in Table 2, these three variables are inde-

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Address reprint requests to Mark L. Williams, PhD, Affiliated Systems Corp., 1200 Post Oak Blvd., Suite 540, Houston, TX 77056-3104. This paper, submitted to the Journal June 26, 1989, was revised and accepted for publication May 16, 1990.

TABLE 1—Seroprevalence Rates for Male IV Drug Users

	Percent Positive/n(%)*	Odds Ratio	95% Confidence Interval (lower/upper)
Race/Ethnicity			
Hispanic	4/107 (4)	—	—
White	14/214 (7)	1.80	.58, 5.62
Black	59/600 (10)	2.80	1.01, 7.90
Frequency of Injection			
<weekly	8/206 (4)	—	—
weekly	25/325 (8)	2.06	.91, 4.67
daily	44/390 (11)	3.15	1.45, 6.82
Length of Injected Drug Use			
0–2 years	2/63 (3)	—	—
3–8 years	26/269 (10)	3.26	.75, 14.13
≥9 years	49/540 (8)	2.77	.66, 11.66
Injection in a Shooting Gallery			
no	22/227 (9)	—	—
yes	55/672 (8)	.92	.55, 1.54
Sharing Needles With a Stranger			
no	16/251 (6)	—	—
yes	61/670 (9)	1.47	.83, 2.60
Number of Persons with Whom Needles Shared			
none	12/139 (9)	—	—
1–2	37/429 (9)	1.00	.51, 1.98
≥3	28/325 (8)	.91	.45, .80
Drug Injected			
heroin	1/32 (2)	—	—
cocaine and heroin	11/198 (6)	1.82	.23, 14.63
cocaine	44/476 (9)	3.16	.42, 23.69
cocaine and amphetamines	21/215 (10)	3.36	.44, 25.85
Syphilis			
no	61/759 (7)	—	—
yes	16/101 (16)	2.34	1.29, 4.24
Gonorrhea			
no	44/562 (8)	—	—
yes	33/359 (9)	1.19	.074, 1.91
Herpes			
no	76/907 (8)	—	—
yes	1/14 (7)	.84	.11, .652
Chlamydia			
no	76/836 (8)	—	—
yes	1/8 (11)	1.38	.17, 11.14
Number of Sexual Partners			
none	9/75 (12)	—	—
one	21/224 (9)	.76	.33, 1.74
2–3	16/220 (7)	.58	.24, 1.36
≥4	31/402 (8)	.61	.28, 1.35

*rounded to nearest whole percent

TABLE 2—Logistic Regression Model

Variables	Odds Ratio	95% Confidence Intervals
Black	1.39	1.05,1.84
Daily Injection	1.40	1.09,1.77
Syphilis	1.42	1.05,1.92

pendently associated with seropositivity. Each of the variables has approximately equal weight.

Discussion

The results of the analysis show that there is an association between race, frequency of injection, and a past history of syphilis. The relationship between frequency of injection and HIV infection did not vary by drug(s) injected. Heroin users injecting daily were just as likely to be HIV positive as cocaine users injecting daily. Although drug injected may have some pharmacological effect on the injecting behaviors of users, how often an IVDU is injecting, whatever the drug, would appear to be of greater importance. The association between a history of syphilis and HIV infection is far from clear. It would seem likely that the variable is indicative of risk of HIV infection due to behaviors associated with the condition.

ACKNOWLEDGMENTS

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HIV Seroconversion in Two Homosexual Men after Receptive Oral Intercourse with Ejaculation: Implications for Counseling Concerning Safe Sexual Practices

ALAN R. LIFSON, MD, MPH, PAUL M. O'MALLEY, NANCY A. HESSOL, MSPH, SUSAN P. BUCHBINDER, MD, LYN CANNON, AND GEORGE W. RUTHERFORD, MD

Abstract: Seroconversion for HIV antibody occurred in two homosexual men who reported no anal intercourse for ≥ 5 years and multiple episodes of receptive oral intercourse with ejaculation. Neither man reported intravenous drug use or receipt of blood products. The last antibody-negative specimen was also negative by the polymerase chain reaction and p24 antigen assays. All sexually active persons should be clearly counselled that receptive oral intercourse with ejaculation carries a potential risk of HIV transmission. (*Am J Public Health* 1991; 81:1509–1511.)

Introduction

Studies of homosexual and bisexual men have shown that anal intercourse, especially receptive anal intercourse with ejaculation, is the sexual practice associated with the greatest risk of human immunodeficiency virus (HIV) infection.[1-5] Although orogenital contact has not been statistically associated in these studies with HIV infection, many subjects engaged in multiple sexual practices, making a relatively lower risk more difficult to detect statistically.

Although transmission of HIV through receptive oral intercourse has been suggested in a few brief reports,[4,6,7] these reports have not provided extensive detail about the circumstances surrounding seroconversion. For this reason, there continues to be controversy about the safety of oral intercourse.

We report seroconversion for HIV antibody in two homosexual men who denied anal intercourse for ≥ 5 years and who engaged in multiple episodes of receptive oral intercourse with ejaculation.

Methods

The San Francisco City Clinic Cohort consists of 6,705 homosexual and bisexual men originally recruited from 1978 through 1980 for studies of hepatitis B.[8,9] Men enrolled in our prospective studies were recruited from this original cohort and are evaluated with an annual interview, including a sexual history by a trained staff member which asks about their activities in the preceding four months and for the entire period since their last interview. Study participants also receive an examination and medical history by a physician unaware of the subject's antibody status. Sera were tested for HIV antibody with the ELISA (Organon Teknika, Research Triangle, NC) and confirmed by both the Western blot and immunofluorescence assay (IFA).[10] Testing for HIV-DNA by the polymerase chain reaction (PCR) was conducted on DNA extracted from peripheral blood mononuclear cells using two different *gag* primers.[11] Sera were also tested for p24 antigen (Coulter Immunology, Hialeah, FL).

Results

For both men who seroconverted, in the two years preceding seroconversion at least three specimens were negative by the ELISA assay; the last pre-seroconversion specimen was also negative by PCR and p24 antigen assays. The first post-seroconversion specimen was positive for HIV antibody by the ELISA, IFA and Western blot assays. Both men were also HIV-antibody positive on repeat testing of the first post-seroconversion specimen (including a second serum sample drawn at the same time), as well as on at least one follow-up specimen drawn between four and nine months later.

Both men reported no insertive or receptive anal intercourse for at least five years preceding the estimated date of seroconversion; this history was supported by three previous interviews. Both men also denied use of intravenous drugs,

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From the AIDS Office, Department of Public Health, San Francisco. Address reprint requests to Alan R. Lifson, MD, MPH, AIDS Office, Department of Public Health, 25 Van Ness Avenue, Suite 500, San Francisco, CA 94102. Dr. Rutherford is currently with the California State Department of Health Services. This paper, submitted to the Journal March 2, 1990, was revised and accepted for publication July 13, 1990.

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The AIDS-Related Experiences and Practices of Primary Care Physicians in Los Angeles: 1984–89

CHARLES E. LEWIS, MD, AND KATHLEEN MONTGOMERY, PHD

Abstract: Telephone interviews of random samples of Los Angeles primary care physicians in 1984, 1986, and 1989 obtained information about their AIDS-related practice experiences, and sexual history taking. Data from mid-1989 reveal almost 74 percent have worked up at least one patient for AIDS or HIV infection in the past six months and 39.5 percent are caring for at least one patient with AIDS or AIDS-related complex. Self-reported use of appropriate sexual history questions has improved substantially over this five-year period. (*Am J Public Health* 1990; 80:1511–1513.)

Introduction

There has been growing speculation, based primarily upon surveys of physicians' attitudes, that a significant proportion of doctors might fail to provide care to patients found to be infected with the human immunodeficiency virus (HIV) and/or who developed acquired immunodeficiency syndrome (AIDS).[1,2]

Over the past six years, we have conducted three telephone surveys of random samples of primary care physicians in full-time office-based practice in Los Angeles, California (January 1984,[3] summer 1985,[4] and summer 1989). Data presented here illustrate changes over this period in physicians' experiences with patients at risk or infected with HIV.

Methods

Stratified random samples of internists, family practitioners, and general practitioners were drawn to reflect the relative proportion of each of these groups practicing in Los Angeles County. The samples were drawn from data tapes provided by the American Medical Association Physician Masterfile, containing listings of all primary care physicians in full-time office-based practice in Los Angeles—a pool of about 5,000 physicians.[5]

Results

Response Rates and Characteristics of the Sample

The response rates for all three surveys have been above 60 percent; the most recent survey resulted in a 77 percent response rate. These response rates are calculated using as the denominator all physicians actually contacted (see Table 1).

Examination of the demographic characteristics of respondents versus nonrespondents during each survey period has revealed no obvious source of bias. The increase in the proportion of women interviewed over the past five years reflects the growing number of women primary care physicians practicing in Los Angeles County; for Los Angeles as

TABLE 1—Response Rates and Characteristics of Respondent Physicians

	1984 (635)[a]	1986 (279)[b]	1989 (463)[a]
% Response Rates	63.1	60.0[c]	77.0
Specialty			
% General Internal Medicine	42.4	44.1	47.7
% General Practice	26.3	17.2	17.3
% Family Practice	31.3	38.7	35.0
% Solo Practice	50.4	54.5	51.2
% Foreign Medical Graduates[d]	17.3	19.9	21.4
% Female	6.8	9.0	19.7

a) Los Angeles County.
b) Los Angeles SMSA.
c) Response rate represents that for the entire state of California (N=1000).
d) Does not include graduates of Canadian or Western European medical schools.

From the Department of Medicine, UCLA School of Medicine (Lewis), and the Graduate School of Management, University of California, Riverside (Montgomery). Address reprint requests to Charles E. Lewis, MD, Division of General Internal Medicine and Health Services Research, B-558 Factor Building, UCLA School of Medicine, 10833 Le Conte Avenue, Los Angeles, CA 90024-1685. This paper, submitted to the Journal January 23, 1990, was revised and accepted for publication April 20, 1990.

a whole, the proportions are: 11 percent in 1984, 14 percent in 1986, and 16 percent in 1989.[5]

AIDS-Related Practice Experiences

The proportion of physicians who, within the six months prior to being interviewed, had worked up or referred a patient for work up of possible HIV infection, including AIDS, has almost tripled, growing from 27 percent to over 73 percent in the five years between 1984 and 1989. Also, the proportion of physicians who have at least one patient in their practice with AIDS has grown from less than 1 percent in 1984 to almost 40 percent in 1989. While 29 percent of primary care physicians surveyed in 1989 had between one and five AIDS patients in their practice, a small group of these physicians (3 percent) were caring for 15 or more patients with AIDS (see Table 2).

Sexual History Taking Questions

Physicians were asked if they took sexual histories on all their new patients. Those responding affirmatively were asked, "What do these questions cover?"

As Table 2 indicates, the proportion of physicians who volunteered they ask about the sexual orientation of their patients has doubled during this five-year interval, with close to half the physicians in 1989 saying they asked for such information. The proportion reporting they ask patients about the number of sexual partners has tripled in the five-year period.

Physicians' Attitudes

Physicians were also asked the extent to which they think the average physician is uncomfortable discussing sexual matters with homosexual patients. In an attempt to avoid only socially acceptable responses, the question was phrased in general terms, rather than asking about the doctor's own level of discomfort. Table 2 shows that these figures have changed little over the past five years, with nearly three-fourths estimating that their colleagues have moderate or considerable discomfort. The estimated discomfort of others is unrelated to physicians' own reported AIDS-related experiences.

Discussion

The data from these three surveys suggest that, as a group, a significant number of primary care physicians in Los Angeles have responded to the growth in the number of AIDS cases by accepting responsibility for the care of patients with AIDS or AIDS-related complex. Unfortunately, these data do not allow us to estimate the extent to which the patients under care of these physicians represent their established patients who have developed AIDS, or whether these are new patients in the practice. Nor can we distinguish among the 60 percent of physicians who have no AIDS patients in their practice how many have refused to care for a patient with AIDS, as contrasted to those who merely have not yet encountered any patients with AIDS.

It may be reassuring to those responsible for funding physicians' education about AIDS that physicians' awareness of the content of a sexual history has improved substantially during the past five years.

While the data on physicians' sexual history taking is self-reported, there is evidence that their awareness of some key questions that should be asked has increased. Knowledge of these is a prerequisite to their use, but we have no data from patients to validate that these were asked. Despite this improvement, a majority of primary care physicians apparently are not asking questions essential in assessing sexual risks.

The persistent belief over time among physicians that their colleagues are uncomfortable in dealing with homosexual patients remains puzzling, given the increases in contact with AIDS patients and reports of more comprehensive sexual histories. Lacking additional data, we were unable to determine whether physicians' perceptions of colleagues' discomfort is unrelated to their own feelings, or that feelings of discomfort simply are not strong predictors of the willing-

TABLE 2—AIDS-Related Experiences and Practices

	1984[a] (95% CI)	1986[b] (95% CI)	1989[a] (95% CI)
Practice Experience			
Worked up a patient for AIDS or HIV infection in past 6 months	27.0 (23.5, 30.4)	43.0 (37.2, 48.8)	74.7 (70.7, 78.7)
Caring for at least one patient with AIDS or ARC	0.6 (−.002, 1.2)	6.1 (3.3, 8.9)	40.0 (35.5, 44.5)
Sexual History Taking[c]			
Ask patients about their sexual orientation (sex of partners)	23.0 (19.7, 26.3)	20.1 (15.4, 24.8)	42.3 (37.8, 46.8)
Ask patients about number of sexual partners	9.9 (7.6, 12.2)	9.0 (5.6, 12.4)	29.8 (25.6, 34.0)
Ask patients about sexual practices	12.1 (9.6, 14.6)	16.5 (12.1, 20.8)	18.4 (14.9, 21.9)
Physicians' Attitudes			
To what extent do you think the average physician is uncomfortable discussing sexual matters with gay patients?			
Considerable	34.9 (31.2, 38.6)	28.5 (23.0, 34.0)	33.3 (28.8, 37.8)
Moderate	41.2 (37.4, 45.0)	43.7 (37.7, 49.7)	38.6 (33.9, 43.2)
Small to None	23.9 (20.6, 27.2)	27.7 (22.3, 33.1)	28.1 (23.8, 32.4)

a) Los Angeles County: 1984 (N = 635); 1989 (N = 463).
b) Los Angeles SMSA: 1986 (N = 279).
c) Percentages are based on the total number of physicians surveyed.

ness of some physicians to provide care in the face of this extraordinary epidemic.

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Behavioral Change in Longitudinal Studies: Adoption of Condom Use by Homosexual/Bisexual Men

Jill G. Joseph, PhD, S. Maurice Adib, MD, MPH, James S. Koopman, MD, and David G. Ostrow, MD, PhD

Abstract: We compared reporting serial cross-sectional prevalence of sexual behavior over time, to reporting individual patterns of behavioral change in a cohort of homosexual men at a six-month interval. Aggregate prevalence rates underestimated the magnitude of change to safer practices, and failed to provide information on relapse to less safe practices. We conclude that it is important to report data based on individual fluctuations in behavior for the evaluation of change over time. (*Am J Public Health* 1990; 80:1513–1514.)

Introduction

During the past few years, the issue of condom use by homosexual/bisexual men in response to the threat of acquired immunodeficiency syndrome (AIDS) has been addressed in several reports from diverse geographic areas. While some reported cross-sectional findings,[1] others have dealt with the more complex task of describing longitudinal behavior patterns.[2] A recent letter by Kelly and St. Lawrence[3] has highlighted the difficulty of documenting relapse to unsafe sex practices through serial cross-sectional studies that mainly monitor relative increases (or decreases) in sexual behavior prevalence.

We have been surveying adoption and maintenance of safer sexual practices, most recently with regard to condom use in receptive anal sex, in a cohort of homosexual/bisexual men in Chicago. This brief report utilizes some findings from that cohort to highlight methodological issues related to reporting behavioral change in longitudinal studies.

Methods

Participants in the Chicago Multicenter AIDS Cohort Study are concurrently enrolled in a longitudinal behavioral and psychosocial investigation. Detailed methods of recruitment and assessment have been described previously.[4,5] Self-administered questionnaires are completed semi-annually after each visit to the study center. Men have already participated in nine such visits and data collecting is ongoing. At visit 4 (V4) in February 1986, 729 men completed a new section of the questionnaire dealing with condom use in receptive anal sex. At visit 5 (V5), complete data were available for 642 men. A total of 601 men completed this section at both visits. Dropout effect was negligible, as participants at V4 and/or V5 did not differ significantly with regard to baseline sociodemographic or behavioral variables. Men in the V4-V5 cohort were predominantly White (89 percent), had a mean age in 1984 of 34.5 years, a mean educational attainment of 16.4 years, and a mean annual income in 1984–85 of $26,300. Forty percent of them were HIV seropositive by V5.

Participants described consistency of condom use during receptive anal sex (always, sometimes, never) with various types of sexual partners, ranging from those with whom they had a steady relationship to anonymous contacts. Two methods for analysis of these data are considered below. The first one defines four levels of condom use in receptive anal sex, from most to least safe.

1. no receptive anal sex;
2. consistent condom use with all partners;
3. inconsistent condom use;
4. no condom use.

This method was applied in serial cross-sectional analyses at V4 and V5. The second method includes these same four components considered in terms of behavioral consistency between V4 and V5. For men in each of the above categories at V4, subsequent behavior at V5 was described in one of three ways:

1. maintenance of the same behavior;
2. change to any less safe practice;
3. change to any safer practice.

Results

Table 1 reports aggregate serial cross-sectional findings from V4 and V5. It shows a slight increase in the prevalence of safer sexual practices (no receptive anal sex or consistent condom use) with some decrease in no condom use. Overall, change reported as aggregate data suggests that prevalence of

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From the Department of Epidemiology, School of Public Health, University of Michigan. Address reprint requests to Jill G. Joseph, PhD, Associate Professor, Department of Epidemiology, School of Public Health, University of Michigan, 109 Observatory, Ann Arbor, MI 48109. This paper, submitted to the Journal March 29, 1989, was revised and accepted for publication April 22, 1990.

TABLE 1—Serial Cross-Sectional Prevalence of Condom Use in Receptive Anal Sex at Visit 4 (V4) and Visit 5 (V5)

	V4 N (%)	V5 N (%)
No receptive anal sex	375 (52)	341 (53)
Consistent condom use	66 (9)	78 (12)
Inconsistent condom use	96 (13)	94 (15)
Condoms never used	192 (26)	129 (20)
Total	729	642

safer sexual practices increased somewhat over a six-month interval.

A second approach to assess risk reduction takes account of intra-individual stability or instability in behavior and is shown in Table 2. Prevalence of intra-individual change in the V4-V5 cohort was 41 percent, much higher than suggested in Table 1. Among those maintaining the same behavior, a majority maintained safer practices, but 49 percent of those never using condoms maintained such a practice over six months. Intra-individual change to safer sexual practices was reported in 24 percent of the cohort, but this approach revealed relapse to less safe sex in 17 percent of participants.

TABLE 2—Individual Patterns of Maintenance and Change in Safer Sexual Practices in Receptive Anal Sex Over a Six-month Period

Practices at V4	Base at V4 N (%)	Maintained at V5 N (%)	Changed to less safe N (%)	Changed to safer N (%)
No receptive anal sex	307 (51)	239 (78)	68 (22)	— —*
Consistent condom use	56 (9)	24 (43)	13 (23)	19 (34)
Inconsistent condom use	85 (14)	27 (32)	23 (27)	35 (41)
Condoms never used	153 (26)	62 (49)	— —*	91 (59)
Total (N = 601)	601	352 (59)	104 (17)	145 (24)

*Such change was not possible at V5, as participants were already at most extreme categories at V4.

Discussion

Intra-individual fluctuations in individual risk-relevant behaviors have been reported in many longitudinal studies of homosexual men.[6] As early as 1985, a report from San Francisco noted that over the course of one year, approximately half the men observed in one cohort had changed their behavior, from completely or probably safe to probably or definitely risky.[7] Aggregate measures, as we demonstrated, can underestimate the magnitude of change to safer sexual practices, but even more important, they often fail to convey important information on adoption, maintenance and relapse in individual patterns of behavior over time. They are therefore, an essential part of evaluating long-term behavioral change, whether the change is related to HIV/AIDS or to other conditions.

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Seat Belt Use in Cars with Air Bags

ALLAN F. WILLIAMS, PHD, JOANN K. WELLS, BS, AND ADRIAN K. LUND, PHD

Abstract: Seat belt use was observed in 1,628 cars with air bags and manual belts and 34,223 cars with manual seat belts only. Sixty-six percent of drivers in cars with air bags wore seat belts compared to 63 percent of drivers in cars with manual belts only. The study found no evidence for the speculation that drivers with air bags will reduce their seat belt use because they believe an air bag alone provides sufficient protection. (*Am J Public Health* 1990; 80:1514–1516.)

Introduction

Federal regulations in the United States require that all passenger cars manufactured in 1990 and after be equipped with automatic restraints. During the phase-in period for this requirement, manufacturers had to equip 10 percent of their passenger cars with automatic restraints in the 1987 model year, 25 percent in 1988, and 40 percent in 1989. Two types of restraint systems have been used to meet the requirement: air bags that inflate automatically in frontal crashes, and seat belts that automatically fasten around occupants when they enter or start the car. When the automatic restraint requirement was being phased in, most manufacturers met it by installing automatic seat belts rather than air bags. Increasingly, air bags have been offered. The percentage of automatic restraint cars sold with air bags was about 10 percent in the 1987 model year and is projected to be 30 percent in

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From the Insurance Institute for Highway Safety. Address reprint requests to Allan F. Williams, PhD, Insurance Institute for Highway Safety, 1005 N. Glebe Road, Arlington, VA 22201. This paper, submitted to the Journal April 23, 1990, was revised and accepted for publication June 29, 1990.

1990 models.[1] Based on announced plans of manufacturers, air bags will be the predominant automatic restraint offered in the 1990s.

Cars with driver side air bags (as well as those few that currently have right front passenger side air bags) are also equipped with manual lap and shoulder belts to meet federal requirements for lateral and rollover crash protection. Together, air bags and lap/shoulder belts provide the most effective restraint protection available. There has been concern that occupants might be less likely to use a manual lap/shoulder belt in vehicles with air bags than in cars that had *only* manual lap/shoulder belts, because they may erroneously assume that the air bag alone provides full or sufficient crash protection. To the extent that this occurred, it would lessen the crash protection otherwise achievable through the combination of air bags and seat belts. To provide information on this question, a survey of belt use in late model cars with and without air bags was undertaken.

Methods

Seat belt use by drivers and right front seat passengers was observed in four metropolitan areas in the fall of 1989. The areas surveyed were metropolitan Chicago, Los Angeles, Philadelphia, and the Maryland and Virginia suburbs of Washington, DC. In each of these areas, seat belt use is required by state law.

To provide a sample of late model cars, observations were restricted to cars with high center mounted brake lights. These lights have been required on all new cars beginning with the 1986 model year. (Some 1985 cars also had high center mounted brake lights, but these were subsequently eliminated from the study sample.)

In each metropolitan area, observations were made at 64 sites, typically intersections with traffic lights. The sites were distributed throughout the metropolitan regions but were concentrated in affluent areas to maximize the likelihood of observing late model cars. Observations at a site were made during either morning (7:00–9:30 am) or evening (4:00–6:30 pm) rush hours. Sites and observation procedures were the same as those used in a previous survey.[2]

In addition to observing shoulder belt use, observers recorded license plate numbers of the vehicles, which were used to obtain vehicle identification numbers (VINs) from motor vehicle departments in Illinois, California, Pennsylvania, Maryland, and Virginia. VINs provide information on car model, model year, and type of restraint system. Using this technique, errors in identifying the vehicle observed are possible only in rare instances, such as the license number was recorded incorrectly or the owner had transferred the license plate from another car to the one observed and the department of motor vehicles had not yet updated their files.

Results

Seat belt use was observed in 45,466 1986–90 model year cars; complete VIN information was available for 97 percent of these vehicles. Of these 44,102 cars, 1,628 were equipped with air bags and lap/shoulder manual belts; 34,223 were equipped with manual lap/shoulder belt systems only. Driver belt use rates were about the same in cars with and without air bags (Table 1). Overall, 66 percent of drivers in cars with air bags used seat belts compared to 63 percent of drivers in cars with manual belts only. Belt use rates of drivers with and without air bags were about the same in each metropolitan area, although belt use rates were substantially higher in Los Angeles and the Washington, DC area than in Chicago and Philadelphia.

TABLE 1—Driver Manual Belt Use in Cars with Air Bags and in Cars with Manual Belts Only, by Metropolitan Area

Metropolitan Area	Percent (N) Using Belts	
	Air Bag Cars	Manual Belt Cars
Chicago	51 (355)	53 (10441)
Los Angeles	73 (684)	71 (8540)
Philadelphia	59 (317)	57 (8521)
Washington, DC	75 (272)	76 (6721)
Total	66 (1628)	63 (34223)

Only a few of the observed cars had a right front passenger air bag and a right front passenger present (N = 22); 50 percent of these passengers were belted. For right front passengers in cars with driver side air bags only, belt use was 56 percent (N = 235) compared to 54 percent (N = 5473) in cars without air bags. Belt use rates for drivers in cars that had right front passengers present was 64 percent in cars with driver side air bags only and 59 percent in cars without air bags.

Table 2 presents driver belt use in cars with and without air bags by manufacturer. Driver belt use was as high or higher in air bag equipped cars as in cars without air bags except for the Chrysler models, which had a slightly lower use rate.

Table 3 presents driver belt use in cars with and without air bags for specific models of cars that were observed in sufficient numbers to make such comparisons. These data indicate that belt use was usually about the same or slightly higher in cars with air bags.

Cars with air bags in the study sample were on average newer than cars with manual belts only. Table 4 presents data on cars with and without air bags by model year. Belt use increased slightly by model year both for cars with and without air bags. For each model year, driver belt use was as high or higher in air bag cars as in cars without air bags.

Discussion

There is no evidence from this survey that the presence of an air bag lessens the likelihood of using a seat belt. Manual seat belt use by drivers of cars with driver side air bags was about the same as for drivers of late model cars that did not have air bags. In cars with drivers and right front passengers

TABLE 2—Driver Manual Belt Use in Cars with Air Bags and in Cars with Manual Belts Only, by Manufacturer

Manufacturer*	Percent (N) Using Belts	
	Air Bag Cars	Manual Belt Cars
BMW	75 (101)	62 (587)
Chrysler	54 (243)	59 (2852)
Ford	64 (157)	59 (5624)
Honda	73 (166)	73 (4186)
Mercedes	64 (772)	—
Volvo	81 (142)	78 (624)
Total	66	64
Total (excluding Mercedes)	67	64

*Forty-seven air bag cars of low volume manufacturers (driver belt use 68 percent) not included.

TABLE 3—Driver Manual Belt Use in Cars with Air Bags and In Cars with Manual Belts Only, by Make/Series

Make/Series	Percent (N) Using Belts	
	Air Bag Cars	Manual Belt Cars
BMW 5,6,7 Series; M6	76 (96)	75 (88)
Chrysler LeBaron Sedan, Convertible	53 (103)	50 (142)
Chrysler LeBaron Coupe	60 (57)	48 (31)
Dodge Daytona	47 (66)	41 (39)
Ford Tempo	57 (46)	59 (282)
Lincoln Continental	68 (90)	66 (77)
Acura Legend	73 (166)	71 (348)
Volvo 740, 760	82 (133)	81 (325)

TABLE 4—Driver Manual Belt Use in Cars with Air Bags and in Cars with Manual Belts Only, by Model Year

Model Year	Percent (N) Using Belts	
	Air Bag Cars	Manual Belt Cars
1986	63 (185)	63 (9781)
1987	65 (319)	63 (9696)
1988	67 (434)	62 (8673)
1989	66 (630)	64 (5994)
1990	67 (60)	66 (79)

present, driver belt use was eight percentage points higher than passenger belt use in cars with driver side air bags and five percentage points higher in cars with manual belts only. If air bags reduced belt use, driver belt use in these air bag equipped cars would have been much lower relative to passenger use.

Manual belt use rates were higher than average in this survey because the observations were restricted to late model cars and were concentrated in affluent areas in states with laws requiring belt use. All of these factors elevate seat belt use.[3,4] Most of the nation's population (89 percent) is now covered by belt use laws; belt use rates in cars with and without air bags would be expected to be lower in the 14 remaining states without laws. Nevertheless, many drivers and right front passengers observed were not using seat belts even though required by law to do so. An important goal is to increase seat belt use both in cars with and without air bags.

ACKNOWLEDGMENTS

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Age, Sex, and Road-Use Patterns of Motor Vehicular Trauma in Rhode Island: A Population-Based Hospital Emergency Department Study

IAN R.H. ROCKETT, PHD, MPH, ELLICE S. LIEBERMAN, MD, DRPH, WILLIAM H. HOLLINSHEAD, MD, MPH, SANDRA L. PUTNAM, PHD, AND HENRY C. THODE, PHD

Abstract: Population-based hospital emergency department data on motor vehicle traffic trauma in Rhode Island, 1984–85, are analyzed by age, sex, and road-use status. Annualized rates of overall and severe trauma were 1,195 cases (95% confidence interval [CI] = 1,164, 1,225) and 102 cases (95% CI = 94, 111) per 100,000 population, respectively. Overall and severe rates peaked at ages 15–24 years. Male rate excesses were most pronounced for motorcycle and pedal cycle trauma. (*Am J Public Health* 1990; 80:1516–1518.)

Introduction

Population-based data on nonfatal motor vehicular trauma are not collected routinely. In population-based studies, case selection has typically been restricted in terms of injury diagnoses, road-use status, and age of population.[1–5] An exception is the Northeastern Ohio Trauma Study, in which a comprehensive sample of motor vehicular trauma cases was drawn from regionwide hospital emergency department records.[6] Adopting a similar methodology, the Rhode Island Department of Health instituted a motor vehicular trauma surveillance system.[7] This report documents patterns of motor vehicle traffic trauma rates for Rhode Island by age, sex, and road-use status.

Methods

Cases represent a 25 percent systematic sample of first encounter motor vehicle traffic trauma patients treated during 1984 and 1985 in the emergency departments of the 12 Rhode Island civilian, acute care, nonpsychiatric hospitals with an emergency facility. They include only Rhode Island patients and fatalities (on RI roads) with known age and sex, who were or could be assigned an ICD-9CM external cause of injury

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From the Department of Health, Leisure, and Safety, and the Health Services Research Center, University of Tennessee, Knoxville; Department of Maternal and Child Health, Harvard School of Public Health, and Department of Obstetrics and Gynecology, Brigham and Women's Hospital, Boston; Division of Family Health, Rhode Island Department of Health; and the Injury Prevention and Analysis Group, Brookhaven National Laboratory. Address reprint requests to Ian R.H. Rockett, PhD, MPH, Department of Health, Leisure, and Safety, University of Tennessee, 1914 Andy Holt Avenue, Knoxville, TN 37996-2700. This paper, submitted to the Journal October 13, 1989, was revised and accepted for publication May 17, 1990.

code (E-code) within ranges E810–E816 and E818–E819. Additional information on the methodology is available elsewhere.[7,8]

Numerators in the overall rates represent annual averages estimated from emergency department cases weighted for the inverse of the sampling fraction. Severe cases were a combination of weighted hospital nonfatal admissions and total fatalities; their initial age groups were collapsed in order to address the small numbers. For both sets of rates, denominators are mid-period age- and sex-stratified population estimates that were based on the 1980 Census for Rhode Island[9] and an official projection of the State population for July 1, 1985.[10]

Results

Overall Trauma

Sample cases numbered 6,268; 92 percent involved Rhode Island residents, and constituted the focus of this report. The estimated 1984–85 average annual number of motor vehicular trauma cases among Rhode Island residents was 11,538, and the corresponding rate was 1,195 cases per 100,000 population (Table 1). Motor vehicle occupants contributed three-fourths of the cases.

The estimated annualized motor vehicular trauma rate for males was 1,311 per 100,000 population, and that for females was 1,080. Irrespective of sex, the rates peaked at ages 15–19 years.* Although varying by sex, this age group also registered the highest rates in the motor vehicle occupant, motorcyclist, and other/unspecified road-use categories. Rates for the pedestrian, pedal cyclist and motorcyclist categories peaked at ages 5–9, 10–14, and 20–24 years, respectively.*

Rates were lowest at the extremes of the age spectrum, and were 21 percent greater for males than females. Male predominance was most evident in the motorcyclist category, where the relative risk exceeded five.

Severe Trauma

The estimated annualized rate of severe trauma (fatalities and hospital admissions) was 103 per 100,000 population (Table 2); 64 percent were in the motor vehicle occupant category; pedestrian and motorcycle trauma cases together comprised almost 30 percent. Respective estimated rates for males and females were 133 and 74 per 100,000 population.

Regardless of sex, the severe trauma rate reached a zenith at ages 15–24 years, and was lowest at ages 0–14 and 45–64. The highest rates in the motor vehicle occupant, motorcyclist and other/unspecified road-use categories were registered in the 15–24 age group. Rates peaked at 0–14 years in the pedal cyclist category. There was a bimodal distribution in the pedestrian category, with rates highest at the extreme ages. Males manifested an 80 percent greater risk of sustaining severe motor vehicular trauma than females. In the motorcyclist category, the overall relative risk for males was nine.

Discussion

Overall trauma rates are probably underestimated here because cases involving individuals who refrained from seeking treatment or were self-treated and those treated in out-of-state hospitals or other health care facilities were outside the study population. Estimated severe trauma rates more closely approximate true rates than do the rates of overall trauma, since they are less likely to be influenced by differential discretionary care-seeking behavior among population sub-groups.

Although the Rhode Island study is more recent than the Ohio study,[6] comparisons are instructive. During the respective time frames of these two studies, neither population was covered by a motorcycle helmet law or by a universal seat belt use law. For every motor vehicle traffic trauma fatality among Rhode Island residents, there were an estimated 11 hospital admissions and 130 patients treated in hospital emergency departments. Respective figures for Northeastern Ohio were one, 11, and 137. Other remarkable similarities are apparent in patterns of overall trauma rates by age, sex, and road-use status as, for example, the excess rates for males and young adults. Severe trauma rates were not computed in the Ohio study.

Experience from the two studies underscores the need for comprehensive E-coding of motor vehicular trauma in hospital records. Also essential in identifying high-risk populations, and designing, targeting, and evaluating appropriate interventions, is compilation and analysis of environmental and behavioral data, such as road configurations at trauma sites and motorist usage of safety devices, alcohol, and other drugs.

Rhode Island was the second state in the nation to enact a child restraint use law. The current data were collected as the founding activity of a nascent injury control unit at the Rhode Island Department of Health with the expectation that they would be useful in advocacy for, and evaluation of, a universal seat belt use law. Unfortunately, Rhode Island remains in the minority of states without such a law. Ironically, absence of a law enabled the Rhode Island

TABLE 1—Number and Percentage of Resident Cases in a 25% Motor Vehicular Trauma Sample by Road-use Status for Persons Treated in Hospital Emergency Departments, and Associated Annualized Rates per 100,000 Population, Rhode Island, 1984–85

	Sample			
Road-use Status	No.	%	Rate	95% Confidence Interval
Motor vehicle occupant	4,319	74.9	894	868, 921
Motorized cyclist	318	5.5	66	59, 73
Pedal cyclist	109	1.9	23	18, 27
Pedestrian	273	4.7	57	50, 63
Other/unspecified	750	13.0	155	144, 166
Total	5,769	100.0	1,195	1,164, 1,225

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*Detailed age-, and sex- and road-use-specific data pertaining to both overall and severe motor vehicular trauma are available from the first author upon request.

TABLE 2—Number and Percentage of Resident Motor Vehicular Trauma Fatalities and Nonfatal Hospital Admissions by Road-use Status, and Associated Annualized Rates per 100,000 Population, Rhode Island, 1984–85

Road-use Status	No. Fatalities*	No. Admissions†	%‡	Rate	95% Confidence Interval
Motor vehicle occupant	84	295	63.8	65	58, 73
Motorized cyclist	18	61	13.2	14	10, 17
Pedal cyclist	4	10	2.2	2	1, 4
Pedestrian	50	67	16.1	16	13, 20
Other/unspecified	21	18	4.7	5	3, 7
Total	177	451	100.0	103	94, 111

*Fatalities represent a complete enumeration of residents who died due to motor vehicular trauma sustained on in-state public roads in 1984 or 1985. The data source is the Division of Vital Statistics, Rhode Island Department of Health.

†Admissions derive from resident cases in a 25% sample for persons treated in hospital emergency departments for motor vehicular trauma. The 35 fatal emergency department cases are excluded.

‡Percentages are based on a combination of total fatalities and weighted nonfatal admissions. The weights are the inverse of the sampling fraction; four.

population to serve as an external control in an ongoing epidemiologic evaluation of the New York State law.[8,11]

The surveillance system which is described here provided data for an alcohol-related injury intervention study.[12] Rhode Island has now established an Office of Injury Control and implemented a variety of interventions aimed at reducing both vehicular and nonvehicular trauma. E-coding by all hospitals is a top priority of the Office, and geographic study of trauma sites is also planned. Both seat belt and helmet use bills were introduced again in the 1990 session of the State Legislature by the Governor.

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Evaluation of the Efficacy and Cost Effectiveness of Health Education Methods to Increase Medication Adherence among Adults with Asthma

Richard A. Windsor, PhD, MPH, William C. Bailey, MD, James M. Richards, Jr., PhD,
Brynda Manzella, MPH, Sen-Jaw Soong, PhD, and Michael Brooks, EdD

Abstract: We randomized 135 adult asthma patients to a control group, and 132 patients to an experimental group which received a special health education intervention. Four adherence measures were documented at baseline and 12-month follow-up: correct inhaler use, inhaler adherence, medication adherence, and total adherence rating. Costs to routinely deliver the intervention were $32.03/patient. Experimental group patients exhibited a significantly higher level of improvement in adherence (44 percent) than control group patients (2 percent). (*Am J Public Health* 1990; 80:1519–1521.)

Introduction

The prevalence of asthma among adults in the United States is 3–6 percent.[1–3] There are over 500,000 hospitalizations and 3,500 deaths from asthma per year.[3] Multiple studies of adults with asthma confirm regimen adherence levels of only 30–40 percent.[4–7] Educational intervention studies have reported methodological problems, used small sample sizes, lacked cost analyses, and presented inconclusive results.[8] This evaluation study was designed to document *efficacy* (the level of behavioral impact produced by a health education program with optimal resources) and *costs* (personnel and materials needed to routinely deliver the program).

Methods

The study was conducted at a university-based comprehensive pulmonary medicine clinic. Patients who had a primary diagnosis of asthma, used medications daily, and met the following diagnostic criteria were eligible: recurrent episodes of wheezing or dyspnea; objective evidence of significantly increased airway resistance during episodes; and improvement in the airway when symptom-free.[9–10] Of the 280 adults $\geq$ 17 years meeting the criteria screened between April 1986 to March 1987, 267 (95 percent) participated.

Following informed consent and baseline assessment, 135 patients were randomized to a control group, and 132 patients to an experimental group. Prior to randomization, patients were stratified by level of asthma severity within each of 11 physician practices (see Appendix). Using pilot study data, an estimated possible improvement in adherence of 20 percent or more for 12 months,[4–8,11–15] and an anticipated 10 percent attrition rate, we defined a need for at least 120 patients in each study group.[16,17]

Health Education Intervention

The experimental group received a peak flow meter and a standardized program from a health education specialist:

- A 30-minute one-to-one session with instruction on peak flow meter use, inhaler use skills, and use of *A Self-Help Guide to Asthma Control;*[18]
- A 60-minute asthma support group session of four to six patients and asthma control partners, and
- Two brief telephone reinforcement calls within one month of the group session.

One-to-one discussions and a review of all methods and materials were conducted with clinic nurses and all 11 participating pulmonary physicians. Four group sessions with patients were held to develop the intervention. All educational components and evaluation methods were pilot tested.

Measurement

All patients received a baseline and 12-month follow-up medical and behavioral assessment including four behavioral outcomes: 1) correct inhaler use, 2) inhaler adherence, 3) medication adherence, and 4) total adherence rating. A 10-item observational checklist documented inhaler use (IU) and skill.[19] Medication adherence (MA) and inhaler adherence (IA) were assessed using instruments adapted from the literature.[20–22] Psychometric analyses (Cronbach Alpha) confirmed adequate reliability: r (MA) = .69, r (IA) = .64, and r (IU) = .78. Items analyses confirmed minimum item discrimination coefficients: $r \geq .20$. A Total Adherence Score was derived by combining patient scores on the medication and inhaler adherence scales.

Over 85 percent of the patients used an inhaled bronchodilator and continuous theophylline. Theophylline levels were analyzed on the two follow-up assessment days to indirectly corroborate patient adherence reports. Patients were informed that their theophylline level would be analyzed to assess adherence. While theophylline level is an imperfect measure of adherence and does not reflect adherence to non-theophylline medications, behavioral studies of adults who know a method is available to validate self-reports document a significant reduction in deception rates (bogus pipeline or pipeline effect).[23,24]

Behavioral Intervention Results

A comparison of the 13 patients who refused (5 percent refusal rate) versus the 267 who participated confirmed no baseline differences by gender, race, age, education, and asthma severity. Data in Table 1 confirm group equivalence at baseline. Thirty-four of the control group patients (25 percent) and eight of the experimental group patients (6 percent) were lost to follow-up. Analyses of baseline data confirmed that the 42 dropouts were not significantly different from the 238 participants. Within the intervention group, use of the *Asthma Guide*[18] was reported by 124 patients (94 percent); 110 (89 percent) participated in the group session

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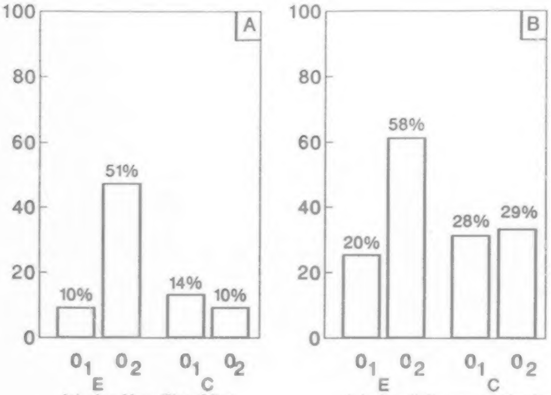
From the Department of Health Behavior, University of Alabama at Birmingham, School of Public Health (Windsor), and the Lung Health Center, UAB School of Medicine, and VA Medical Center (Bailey, Richards, Manzella, Soong, Brooks). Address reprint requests to Richard A. Windsor, PhD, MPH, Chair, Department of Health Behavior, UAT School of Public Health, UAB Station—CBB 502, Birmingham, AL 35294. This paper, submitted to the Journal June 26, 1989, was revised and accepted for publication May 1, 1990.

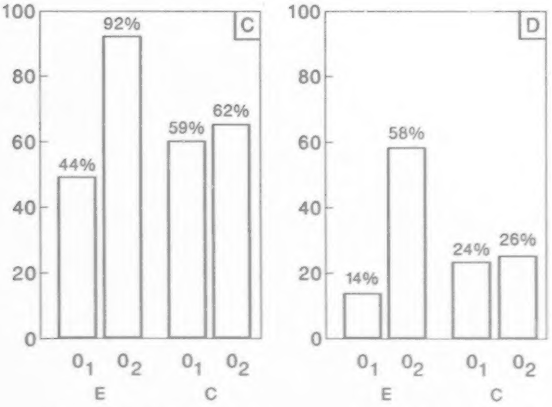
TABLE 1—Base-Line Patient Characteristics by Study Group

Characteristics	Control Group N = 135	Experimental Group N = 132
% Female	71	6[illegible]
% Black	28	32
Median Age (years)	49	50
Median Years of Education	13	13
% Current Smoker	13	10
% Asthma Severity		
Mild	39	37
Moderate	45	48
Severe	17	16



Inhaler Use Checklist
All Ten Items - Correct

Inhaler Adherence Scale
All Six Items - Yes



Medication Adherence Scale
All Six Items - Yes

Total Adherence Score
All Twelve Items - Yes

FIGURE 1—Behavioral Impact by Type and Study Group

and 124 patients (94 percent) received both reinforcement calls.

The behavioral impact of the intervention is presented in Figure 1. Using a 95% confidence interval (CI) to evaluate differences between rates of group improvement, a consistent pattern of adherence was confirmed for a 12-month period. Significant improvements for the intervention group in inhaler skills use (CI = .29, .61), inhaler adherence (CI = .24, .50), medication adherence (CI = .31, .57), and total adherence score 100 percent adherence (CI = .28, .56) were observed. Little behavior change was observed in the control group.

Cost Estimation and Analysis

Patient time and intervention development costs were not used to compute program costs. While an MPH-health education specialist provided the intervention, a nurse would be the typical provider. We used a salary of $25,000 plus a fringe benefit rate of 20 percent to estimate nursing cost: $30,000 per year/2080 hr = $14.42/hr. Total personnel time costs were $24.30: Component #1 = $7.21, Component #2 = $14.42 and Component #3 = $2.40. The *Asthma Guide*[18] cost $8.00 for a Total Intervention Cost = $32.03/patient. The total control group costs for a 10-minute general discussion about the importance of adherence and two brief follow-up contacts for a total of 15 minutes were $3.61/patient.[25,26] Cost analyses presented in Table 2 confirmed the need to allocate resources to significantly increase patient adherence.

Discussion

The increases in adherence observed in our intervention group exceeded those of previous reports. While some proportion of patient self-reports of adherence may represent a "social desirability" response, psychometric analyses indicated good measurement of behavioral impact. Our high recruitment rate confirmed that adults with asthma are interested in taking an active role in their self-management and will participate in a health education program integrated into routine asthma care. Near unanimous positive anecdotal comments about the intervention by nurses, physicians, and patients were observed. Feasibility, patient and provider acceptance was very high. From an administrative and programmatic perspective, our experience indicated that this type of intervention has potential for adaptation by other asthma programs. However, in our attempt to maximize efficacy, an intervention was provided which may be too resource intensive and costly for other asthma care sites.

While the methods and results reported here and elsewhere were encouraging,[8,18,27] future health education research needs to evaluate the impact of a streamlined health education program on health services utilization, e.g., am-

TABLE 2—Cost Effectiveness of the Health Education Program

Groups	Total Cost per Patient	Total Adherence Score Improvement	Cost Effectiveness*
Control Group (N = 135)	$ 3.61	+2%	$243.68
Experimental Group (N = 132)	$32.03	+44%	$ 96.09

*Cost Effectiveness = Total cost per group ÷ Adherence score increase

bulatory care, emergency care and hospitalization, and clinical outcomes.

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The findings were presented at the 117th annual meeting, American Public Health Association, October 25, 1989, Chicago, Illinois.

APPENDIX

Dear Doctor,

As part of the Asthma Education Project, the severity of your patient's asthma based on their current regimen is needed.

Please circle the appropriate severity level:

0—Mild—generally wheeze-free, responds well to medication.
1—Moderate—occasional wheezing with periods of significant wheezing.
2—Severe—constant wheezing, does not respond well to medications, steroid dependent

Please encourage your patient to participate in the Asthma Education Project.

Thank you,

The Asthma Project Staff

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Findings on Health Status and Health Care Needs

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HISPANIC HEALTH AND NUTRITION EXAMINATION SURVEY, 1982–84:

Findings on Health Status and Health Care Needs

FERNANDO M. TREVIÑO, PhD, MPH
EDITOR

ADOLPH P. FALCÓN, MPP
AND
CHRISTINE A. STROUP-BENHAM, MA,
ASSISTANT EDITORS

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The National Coalition of Hispanic Health and Human Services Organizations (COSSMHO) is located at 1030 15th Street, NW, Suite 1053, Washington, DC 20005, Telephone: (202) 371-2100.

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HISPANIC HEALTH AND NUTRITION EXAMINATION SURVEY, 1982–84:

Findings on Health Status and Health Care Needs

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FOREWORD

In 1986 the National Coalition of Hispanic Health and Human Services Organizations (COSSMHO) established the Hispanic Health Research Consortium (HHRC) to spearhead analysis of the landmark Hispanic Health and Nutrition Examination Survey (HHANES). Conducted by the National Center for Health Statistics (NCHS), HHANES was the first representative survey to provide physical examination information on the health and nutritional status of Mexican Americans, Puerto Ricans, and Cuban Americans.

HHRC was established when COSSMHO entered into an agreement with the Office of the Assistant Secretary for Health (OASH) in the US Department of Health and Human Services and the National Center for Health Statistics to facilitate research by nongovernment sources on the data from the HHANES. Two rounds of competitive grants were awarded by COSSMHO to establish a network of eight university and community-based research teams. COSSMHO coordinated the research of this network, provided technical assistance, convened meetings of HHRC researchers, and disseminated information nationally to researchers interested in conducting HHANES analyses.

The unique partnership between COSSMHO, OASH, NCHS, and university and community-based researchers ensured that there would be: support for HHANES research in critical areas; appropriate outreach to Hispanic researchers and those who have developed a reputation for their work on health in Hispanic communities; support for graduate level students to conduct research on the health status of Hispanic communities; and hopefully the development of a pool of future research talent. The establishment of the Hispanic Health Research Consortium has moved the analysis of Hispanic health care issues to a new level of sophistication.

The information presented in this supplement and available for HHANES represents a significant increase in our nation's ability to examine and understand the health of Hispanic Americans. It is our hope that this publication and subsequent analyses of HHANES data will provoke thought and debate and will serve as an impetus for new research on the health status and health care needs of our nation's 23 million Hispanic men, women, and children.

JANE L. DELGADO, PhD, MS
President and Chief Executive Officer
National Coalition of Hispanic Health and Human Services Organizations (COSSMHO)

ROBERT S. MURPHY
Director, Division of Health Examination Surveys
National Center for Health Statistics, CDC

CONTRIBUTORS

Hortensia Amaro, PhD, Associate Professor, School of Public Health, Boston University School of Medicine, Boston, Massachusetts

Gerald Coffman, MS, Data Analyst, School of Public Health, Boston University School of Medicine, Boston, Massachusetts

Bernice H. Cohen, PhD, Department of Epidemiology, School of Hygiene and Public Health, Johns Hopkins University, Baltimore, Maryland

Jane L. Delgado, PhD, MS, President and CEO, National Coalition of Hispanic Health and Human Services Organizations (COSSMHO), Washington, DC

Brenda Eskenazi, PhD, Faculty Member, Maternal and Child Health Program and Epidemiology Program, University of California, Berkeley, California

Antonio L. Estrada, PhD, MSPH, Associate Director and Assistant Professor, Southwest Border Rural Health Research Center, College of Medicine, Department of Family and Community Medicine, University of Arizona, Tucson, Arizona

Melinda Garcia, PhD, Assistant Professor, California School of Professional Psychology, Los Angeles, California, and Clinical Director, Substance Abuse Treatment Center, El Centro, California

Jeffrey B. Gould, MD, MPH, Faculty Member, Maternal and Child Health Program, University of California, Berkeley, California

Sylvia Guendelman, DSW, Faculty Member, Maternal and Child Health Program, University of California, Berkeley, California

Clair Harvey, Smoking, Tobacco, and Cancer Program, National Cancer Institute, Bethesda, Maryland

Suzanne G. Haynes, PhD, Health Promotion Sciences Branch, National Cancer Institute, Bethesda, Maryland

Tim Heeren, PhD, Associate Professor, School of Public Health, Boston University School of Medicine, Boston, Massachusetts

John C. Higginbotham, MPH, Division of Sociomedical Sciences, Department of Preventive Medicine and Community Health, The University of Texas Medical Branch, Galveston, Texas

Mark Hudes, PhD, Senior Statistician, Department of Nutritional Sciences, University of California, Berkeley, California

Amid I. Ismail, BDS, MPH, DrPH, Division of Community Dentistry, Faculty of Dentistry, Dalhousie University, Halifax, Nova Scotia, CANADA

Clifford L. Johnson, MSPH, Chief, Nutrition Statistics Branch, National Center for Health Statistics, Hyattsville, Maryland

Kyriakos S. Markides, PhD, Professor, Department of Preventive Medicine and Community Health, University of Texas Medical Branch, Galveston, Texas

Gary Marks, PhD, Assistant Professor, Department of Preventive Medicine, University of Southern California, Los Angeles, California

Henry Montes, Henry J. Kaiser Family Foundation, Menlo Park, California

Robert S. Murphy, Director, Division of Health Examination Surveys, National Center for Health Statistics, 6525 Belcrest Road, Hyattsville, MD 20782

Herbert Nickens, MD, Association of American Medical Colleges, Washington, DC

Laura A. Ray, MPA, Programmer Analyst, Center for Cross-Cultural Research, University of Texas Medical Branch, Galveston, Texas

Ila Roy, MSW, Former Research Assistant at The National Coalition of Hispanic Health and Human Services Organizations (COSSMHO), Washington, DC

David Shelton, MS, Data Coordinator, Division of Hematology and Oncology, Children's Hospital, Los Angeles, California

Julia M. Solis, PhD, Senior Research Associate, Comprehensive Cancer Center, University of Southern California, and current Director, Local Policy Studies, Tomas Rivera Center, Claremont, California

Christine A. Stroup-Benham, MA, Associate in Medical Education, Office of Educational Development, University of Texas Medical Branch, Galveston, Texas

Susan M. Szpunar, RDH, MPH, DrPH, Program in Dental Public Health, School of Public Health, The University of Michigan, Ann Arbor, Michigan

Dorothy B. Treviño, PhD, Department of Family Medicine, University of Texas Medical Branch, Galveston, Texas

Fernando M. Treviño, PhD, MPH, Director, Center for Cross-Cultural Research, University of Texas Medical Branch, Galveston, Texas

Rupert Whitaker, Research Fellow, School of Public Health, Boston University School of Medicine, Boston, Massachusetts

---

Adolph P. Falcón, MPP (not a contributing author, is assistant editor of the supplement); his primary affiliation is: Senior Policy Advisor, National Coalition of Hispanic Health & Human Services Organizations, Washington, DC.

I. Hispanic Health and Nutrition Examination Survey: Methodological Considerations

Jane L. Delgado, PhD, Clifford L. Johnson, MSPH, Ila Roy, MSW, and Fernando M. Treviño, PhD, MPH

Abstract: The Hispanic Health and Nutrition Examination Survey (HHANES) was the first special population survey undertaken by the National Center for Health Statistics. The HHANES was designed to assess the health and nutritional status and needs of Mexican Americans, mainland Puerto Ricans and Cuban Americans. Data were collected using five data collection techniques: direct physical examinations, diagnostic testing, anthropometry, laboratory analyses, and interviews. Unlike other surveys conducted by the National Center for Health Statistics, the HHANES was not designed as a national survey. The HHANES was a survey of three Hispanic subgroups of the population in selected areas of the United States with a survey universe that included approximately 76 percent of the 1980 Hispanic-origin population in the United States. This article discusses statistical issues that should be addressed by researchers when analyzing HHANES data. Specifically, analysts need to account for the complex sample design, nonresponse bias, potential non-coverage bias, and the regional nature of the HHANES sample. [*Am J Public Health* 1990; 80(Suppl):6–10.]

Introduction

HHANES is the largest and most comprehensive Hispanic health survey ever carried out in the United States. It collected a wide spectrum of socioeconomic, health and demographic data; administered physical and dental examinations; took diagnostic specimens and conducted specialized tests, producing one of the most complete and unique Hispanic health data bases available to date.

History

The roots of the survey can be traced to 1977 when the National Academy of Public Administration was contracted to evaluate the National Health and Nutrition Examination Survey (NHANES). An expert panel, formed to assess NHANES utility in addressing nutrition-related public health issues, recommended that the National Center for Health Statistics (NCHS), the agency of the Public Health Service commissioned to carry out NHANES, should consider studies of the health and nutrition status of major subpopulation groups which could yield estimates comparable to those of the general US population.[1] The Hispanic population was subsequently selected as the special population group to be surveyed due to its size and the lack of available health status data. This study enables researchers to compare results available for the general population.

Planning for the survey began in 1979 with actual data collection occurring over a two-year period (1982–84) in three regions of the United States. HHANES surveyed Hispanics from the three major Hispanic subgroups as indicated by the 1980 Census: Mexican Americans, Puerto Ricans and Cuban Americans. Analysis of HHANES data began in earnest in 1986 when agreements were entered into with the National Coalition of Hispanic Health and Human Services Organizations (COSSMHO) to establish the Hispanic Health Research Consortium (HHRC).

Design

HHANES became the first special population survey undertaken by the National Center for Health Statistics to assess the health and nutritional status and needs of Hispanics and to investigate the access to services by this population. HHANES was designed to focus on certain selected chronic conditions rather than to provide a comprehensive picture of health status or total need for care.[2] It was also intended to provide baseline health and nutritional information for Hispanics comparable to other surveys conducted by NCHS, particularly the Second National Health and Nutrition Examination Survey (NHANES II).[3]

The stated goals of HHANES were: to conduct a study of the health and nutritional status of the Hispanic population; to produce information useful for public health planners and researchers; and to stimulate further research and investigations on Hispanic health issues. The types of information that health examination surveys contribute are related to specific conditions or measures and include: total prevalence of a condition or characteristic, descriptive or normative information, and relationships between health variables and demographic factors. HHANES was designed to enable the study of all of these aspects of health.

The major chronic conditions studied by HHANES were: diabetes, hypertension, heart disease, gallstones, dental disease, otitis media and hearing problems, kidney disease, liver function, and depression. In addition, it surveyed the prevalence of immunization, alcohol consumption, cigarette consumption, and drug abuse in the population. HHANES also had as a major focus nutritional status, nutrition-related conditions (iron status and anemia, Vitamin A levels), and food consumption patterns as related to overweight, diabetes, alcohol consumption, dental health, and digestive diseases. Environmental exposure was a final aspect of the survey that measured exposure by the sample population to lead and selected pesticides.

Apart from these medical issues, questionnaires included numerous demographic and economic items. Income assistance, insurance coverage, health services use and satisfaction, acculturation, and assessment of health (both self-assessment and physician assessment) are examples of information collected on the various questionnaires for each sample person.

The Acculturation Scale

Many of these subjects are critical factors in the analyses of HHANES data, but pose special problems of their own. For example, the complex nature of acculturation makes it difficult to sort out this factor as a mediator of health problems. The eight variable acculturation scale used in the HHANES was modified from a 20-item scale developed for

Note: Author affiliations are listed elsewhere under CONTRIBUTORS.

and validated with a Mexican origin population.[4] It was not designed to assess acculturation levels of Puerto Ricans or Cuban Americans. The eight item HHANES acculturation scale assessed language ability, self-identification, parents' ethnic identification and generation in the US. Sample persons were asked: 1) what language do you speak? 2) What language do you prefer? 3) Can you read Spanish? English? Which do you read better? 4) Can you write Spanish? English? Which do you write better? 5) How do you identify yourself? 6) What ethnic identification does/did your mother use? 7) What ethnic identification does/did your father use? and 8) Where were you born? Your mother? Your father? Each of the first seven variables is scaled from 1 to 5, where 1 represents the least acculturated category and 5 represents the most acculturated category. Variable 8 (birthplace) cannot be scaled from 1 to 5, as it was in the original scale, because of gaps in the data collected in the HHANES. In the original scale developed by Cuellar, *et al*, a first generation American was scored as a "1" (born in Mexico) and a fifth generation person was scored as a "5". The HHANES did not collect information regarding grandparents' birthplace, thus persons could only be classified as being first, second, and third or higher generation Americans. Persons who were not born in Mexico and whose parents were also not born in Mexico were coded as a "4" (third or higher generation). Nonetheless, the detailed nature of the instruments developed and administered by the HHANES has the great potential to provide a uniquely rich picture of various aspects of illness, health, and health service use (as well as barriers) in the Hispanic population.

The Target Population

It is important to note that HHANES was not designed as a national Hispanic survey; therefore, no national estimates for Hispanics can be drawn based on HHANES findings. The major difference between HHANES and previous NHANES surveys is that HHANES was a survey of three Hispanic subgroups of the population in selected areas of the United States rather than a national probability sample. Even though HHANES was not designed to be representative of all Hispanics residing in the United States, the survey universe included approximately 76 percent of the 1980 Hispanic-origin population in the United States. The three Hispanic subgroups and the area covered were: Mexican Americans residing in five southwestern states (Arizona, California, Colorado, New Mexico and Texas); Cuban-Americans residing in Dade County, Florida; and Puerto Ricans residing in the New York City metropolitan area including parts of New York, New Jersey, and Connecticut.

The target population for HHANES ideally would have been all households in the United States with at least one member of Hispanic origin. However, there are many states and counties with very small numbers or proportions of Hispanic persons. Therefore, HHANES was restricted to those areas of the country described above in order to permit an efficient operation of the survey from the standpoints of cost and logistics. Even within the areas defined for the survey, there were certain counties excluded from the survey design for the same reasons. This may result in some small bias due to noncoverage. The sample weights created for analysis and estimation purposes attempt to take into account this issue of noncoverage.

Sampling and Respondent Selection

Survey results can be generalized to these subgroups and geographic areas. The design included identification of three universes: a Mexican-origin universe made up of 193 primary sampling units (PSUs) in five Southwestern States (Texas, Colorado, New Mexico, Arizona and California); a Cuban-origin universe consisting of one PSU (Dade County, Florida); and a Puerto Rican-origin universe consisting of 16 PSUs from New York, New Jersey, and Connecticut. The PSUs were stratified to group similar counties and one was selected from each stratum. Finally, segments and households were selected from each PSU.

Within each PSU, block areas with less than a minimum number of eligible Hispanics were excluded due to the costs of interviewing ineligible households. This exclusion should be considered during interpretation of the results. Households which had at least one Hispanic member (for example, a Mexican American in the Southwest) were eligible to participate; therefore it should be noted that a small percent of non-Hispanics are included in the overall sample.

Selected households were screened to identify eligible Hispanic families and to select sample persons from these families to be interviewed and examined. Eligibility for the survey was determined by the family unit. A family was considered to be eligible if at least one family member's reported national origin or ancestry met the criteria for eligibility appropriate to the survey location. In cases where multiple origins were reported for the same person on different questionnaires in the survey, the person was considered eligible if any one of the reported origins met these criteria.

If a family was eligible for the survey, then every member of that family was eligible to be selected for the interview and examination procedures. Therefore, some non-Hispanic persons residing in Hispanic households and some Hispanic persons not meeting the above criteria were selected for the survey and examined. Thus, analysis of HHANES data can be done in one of two ways. If all sample persons are included, then the analysis is of Hispanic households (which may include non-Hispanic individuals). If only sample persons meeting the eligibility criteria are included, then the analysis is of Mexican Americans, Cuban Americans, or Puerto Ricans, respectively.

Data Collection

HHANES relied on five data collection techniques: direct physical examinations, diagnostic testing, anthropometry (body measurements), laboratory analyses, and interviews. Depending on the nature of the interviews, some were administered in the household and others were administered in private within a mobile examination center. All examinations and testing were performed in the mobile examination center located in the community.

Due to the survey's modus operandi, attrition occurred both at the in-house interviews and at a later examination in the mobile trailer. The attrition of sample persons to the health examination occurred for a variety of reasons, e.g. had undergone a recent examination, fear of having an examination, lack of time, etc. Therefore, there are interview data on a larger sample than examination data.

Oral glucose tolerance tests and an ultrasound examination of the gall bladder were administered to a half-sample of the adult population; these people were asked to fast overnight and were examined in the morning sessions.

Caution should be exercised when utilizing data from certain specialized examinations. Due to the intrusive and demanding nature of the tests, or special requirements placed to ensure accurate testing, a low response rate resulted, particularly in the case of the glucose tolerance test.

Virtually all interview staff involved in the survey were bilingual; most were also bicultural.

Response Rates

NCHS has estimated that the HHANES survey represented approximately 76 percent of the total 1980 US Hispanic civilian, noninstitutionalized population aged 6 months through 74 years. Table 1 shows the sample sizes and response rates for each of the three survey areas in HHANES. The results are presented for both the total sample and the specific origin sample as explained previously.

Statistical Guidelines

Each of the authors of the articles in this supplement was provided with detailed statistical guidelines. Because the Health and Nutrition Examination Surveys have used a complex survey design rather than a simple random sample design, standard statistical methods may not be appropriate for data analysis. Guidelines for NHANES data analysis which produce the appropriate covariance structure and population estimates have been described elsewhere, and will only be summarized here.[2,5] Sample weights must be used to produce correct population estimates since each sample person does not have an equal probability of selection. The weights incorporate the selection probabilities, and adjustments for post-stratification (in the Mexican American sample only), noncoverage and nonresponse. In addition to the sample weights, the strata and the primary sampling units from the sample design are needed to estimate the sample variances and to test for statistical significance. In general, sampling variances will be underestimated if calculated without incorporating the complex sample design. (*Note:* the complex sample estimate of sampling variance includes variability due to measurement error).

The effect of the complex sample design on variance estimates is reflected by the design effect, which is the ratio of the variance of a statistic from a complex sample to the variance of the same statistic from a simple random sample of the same size.[2] A design effect of one indicates the equality of the simple random sample variance and the complex sample variance.

TABLE 1—Sample Size and Response Rates by Survey Area and Specified Hispanic Origin: Hispanic Health and Nutrition Examination Survey, 1982–84

Survey Area and Hispanic Origin	Sample Size	Interviewed		Examined	
		No.	Percent	No.	Percent
Southwest area					
All Persons	9,894	8,554	86.5	7,462	75.4
Mexican American	9,455	8,222	87.0	7,197	76.1
Dade County, Florida area					
All Persons	2,244	1,766	78.7	1,357	60.5
Cuban	2,125	1,677	78.9	1,291	60.8
New York City area					
All Persons	3,786	3,369	89.0	2,834	74.9
Puerto Rican	3,525	3,137	89.0	2,645	75.0

However, design effects in NHANES I and II were rarely one, and the magnitude of the design effects was highly variable. Although preliminary analyses may be performed on unweighted data with standard statistical packages that assume simple random sampling, final analyses should be done on weighted data using special computer programs that use a balanced half-sample replication methods (e.g. SESUDAAN, SURREGR) which incorporate the complex sample design.[6,7]

The complex survey design with effects greater than 1 also increase the minimum sample size needed for calculation of reliable estimates of statistical parameters. The sample size requirements were calculated assuming an average design effect of 1.5. The approach was also used in the initial analyses of the HHANES data, because HHANES also used a complex sample design. However, analyses performed during data editing revealed a number of unexpectedly large (>7.5) or small (<.8) design effects.[8] The unstable variance estimates did not show a consistent pattern, nor were they alleviated by using different methods of estimating variances, stratifying, or incorporating the third level of sampling. In general, then, calculating point estimates of sample variances by using the special computer programs described earlier cannot be recommended. An interim recommendation to use average design effects to correct variances based on the assumption of simple random sampling was made for HHANES data analysis. This is not the only approach that could be used to deal with this issue. It is, however, a relatively simple and practical solution. The impact of using this approach for various types of analyses is discussed in following chapters.

Average design effects have since been calculated for many analytic variables for the three Hispanic subgroups. In general, the average design effects for the Mexican American and Puerto Rican subgroups tend to average 1.5, while those for the Cuban subgroup tend to be about 1. Thus, in developing preliminary guidelines for analysis and presentation of HHANES data, these differences in design should be considered.

Criteria for Presentation of Data

Using the information described above about the effects of the HHANES complex survey design, the following guidelines were developed for the reporting of means and percents for the HHANES data:

- If the sample size in an analytic cell is less than 25, the mean or percent should not be reported;
- If the sample size is 25–44, the mean or percent should be reported with an asterisk beside it to indicate that the statistic does not meet a minimum reliability standard;
- If the sample size is 45 or more, the mean or percent is presented without caveat.

Exceptions to these guidelines are possible in selected instances in order to more completely present a series of findings, as long as the appropriate caveats are clearly indicated.

Multivariate Analyses

Multivariate analyses should be performed in three stages as recommended by Landis, *et al*,[5] and Kovar.[8] The first stage involves running the multivariate analyses without regard to the weights or complex sampling design. The second stage consists of running the same analyses incorporating the sample weights. The third and final stage involves running the same analyses incorporating the sample weights and the complex survey design.

The second and third stages of analyses can use a transformation or "Standardization" of the statistical weight variable to arrive at unbiased test statistics. To normalize the weights to the actual sample size of the analytic sample the following equation is used:

> Normal Weight per sample person = (individual weight) * (actual sample size of analytic sample/the sum of the weights of the analytic sample).

Analysts should use software packages or procedures in stage three analyses that incorporate the complex survey design (e.g. SURREGR and SUPER CARP).[7,9] If this is not possible, other standard software packages or procedures (e.g. SAS PRICER) can be used.[10] In this latter situation, using the normalized weights and exercising caution in interpreting test statistics is essential. For example, analysts need to consider whether a significant t-statistic for a Beta coefficient would remain significant with a design effect of 2.

This can be accomplished by multiplying the denominator of the t-statistic by the square root of 2. If the t-statistic would remain significant with a survey design effect of two (or even higher), then the analyst could feel fairly comfortable with the result. However, if the result would change with a design effect somewhere between one and two, then the interpretation of the finding should be qualified and supported by other findings in the literature.

Considerations of reliability of test statistics also require the stage three analyses incorporate some allowance for the variability as well as the size of the design effect. Analysts were advised to limit the number of independent variables per analysis to no more than the number of degrees of freedom for a given analysis, usually defined as the number of PSUs minus the number of strata (see B. V. Shah's discussion in the documentation for SURREGR). For example, separate regression analyses for NHANES II Whites and Blacks should include no more than 32 and 16 variables, respectively; and analyses of Mexican Americans, Cuban Americans, and Puerto Ricans should include no more than eight variables. For the same reason, (i.e. considerations of reliability), the forward approach is preferred to the backward approach in variable selection for regression analyses.

In summary, multivariate analyses should always be undertaken with a purpose in mind and caution in interpretation of the results. The complex NHANES and HHANES surveys with relatively small numbers of PSUs, and, in some cases, sample sizes, require the analyst to exercise good judgment and to realize that some of their analyses would result in more hypothesis-generating than answering.

Comparing NHANES and HHANES Results

The average design effect approach can be used to calculate complex sample variances for both NHANES data and HHANES data as described previously. This results in one common approach being used to handle the complex sample design for both surveys. A second approach involves calculating individual complex variance estimates from the appropriate software programs mentioned previously for all races and for Whites from NHANES I or NHANES II. These are combined with average design effect variance estimates for Blacks from NHANES I or NHANES II and Mexican Americans, Cuban Americans, and Puerto Ricans from HHANES. This mixed approach is much more difficult to explain to the reader, but is an acceptable alternative. Regardless of the approach selected, a t-statistic should be used in any statistical testing of differences between the various population groups.

Conclusions

It is important, apart from its specific scope, to be aware of the limitations of HHANES data in terms of the manner in which the sample design and response rates affect analyses and interpretation of the findings. The interpretation of results should be made with the following factors in mind:

- *HHANES does not have a simple random sample design.* Only weighted estimates are considered accurate in reflecting the sub-population to which the sample person participating in HHANES belongs. It is necessary, during analysis, to incorporate the sample weights provided by NCHS on each data tape. Because of the complex sample design and the ratio adjustments used to produce the sample weights, commonly used methods of point and variance estimation and hypothesis testing which assume simple random sampling may give misleading results.
- *Nonresponse bias in HHANES.* There exists the potential for three levels of nonresponse: household interview nonresponse, examination nonresponse, and item nonresponse. The potential effect of any nonresponse bias is greater when response rates are low. Therefore, intense efforts were undertaken during HHANES to develop and implement procedures and inducements that would reduce all types of nonresponse and thereby reduce the potential for bias in the survey estimates. Despite these efforts, a number of sample persons did not participate in the survey (see Table I). Therefore, it is important that any analysis of HHANES data include some assessment of the possible impact of nonresponse bias on the results.
- *Potential non-coverage bias in HHANES sample.* Due to the exclusion of areas with few Hispanic residents, there is a potential non-coverage bias in the HHANES sample that should be considered when interpreting results.
- *HHANES was not designed as a national Hispanic survey.* No national estimates for Hispanics can be drawn based on HHANES findings. Neither can the data for the three Hispanic subgroups be combined to produce a single estimate for the entire Hispanic population. Only regional estimates for the three subpopulations included are possible.

Due to its complex nature, it is important to be thoroughly familiar with the HHANES data base for its applicability to research areas under consideration and for analytical procedures suggested. This is facilitated by ample documentation now available on the subject as well as the documentation accompanying each data tape. When properly used, the HHANES data base constitutes an invaluable and unique resource for studying the health and nutritional status of the Hispanic population.

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II. Acculturation, Access to Care, and Use of Preventive Services by Hispanics: Findings from HHANES 1982–84

JULIA M. SOLIS, PHD, GARY MARKS, PHD, MELINDA GARCIA, PHD, AND DAVID SHELTON, MS

Abstract: Use of preventive health services (physical, dental, and eye examinations, Pap smear and breast examinations) among Mexican American, Cuban American, and Puerto Rican adults (ages 20–74) was investigated with data from the HHANES. Analyses focused on the relative importance of two predictors of recency of screening: access to services (health insurance coverage, having a routine place for care, type of facility used, having a regular provider, travel time) and acculturation (spoken and written language, ethnic identification). Regression analyses controlling for age, education, and income indicated that utilization of the preventive services was predicted more strongly by access to care than by acculturation. For each Hispanic group, having a routine place for health care, health insurance coverage, and a regular provider were each significantly associated with greater recency of screening. Type of facility used and travel time produced less consistent effects. These results replicate past studies that have demonstrated the important link between institutional access and use of health services. Of the acculturation variables, language but not ethnic identification (which was measured only for the Mexican Americans) predicted use. This latter finding, which has been demonstrated in other studies as well, suggests that the effect of language on screening practices should not be interpreted as a cultural factor, but as an access factor, i.e. use of English favors access to services. [*Am J Public Health* 1990; 80 (Suppl):11–19.]

Introduction

Despite the steady improvement in the health status of Americans generally, Hispanics in the United States continue to be at greater risk for health problems.[1] This risk may stem, in part, from underutilization of screening services for the prevention and early detection of disease. Mexican Americans, for example, are less likely than non-Hispanic Whites to use such services as routine medical check-ups, dental and eye examinations, prenatal care, and family planning.[2–6]

Two general hypotheses can be invoked to account for Hispanics' underutilization of preventive health services. One hypothesis argues that level of acculturation of Hispanics influences their utilization patterns.[7–9] As discussed by Keefe,[10] acculturation is the process of change that occurs as a result of continuous contact between cultural groups. Although such a broad definition does not specify the types, degree, or direction of change expected, studies have largely presumed that Hispanics who have adopted the behavioral practices and values of the dominant society are more likely to utilize health services.[8]

Indirect support for this view comes from studies indicating that Mexican Americans used preventive services less than the general population,[3–5,11] even after controlling for sociodemographics and health insurance coverage. These studies, however, did not measure acculturation directly; thus they can only presume that the differences are due to cultural factors. Other studies have measured acculturation with multidimensional scales or with measures of language preference. Some suggest an effect of acculturation on utilization of health services,[6,12–14] whereas others fail to find an effect.[15–17] Despite these inconsistent results, a trend is apparent. Those studies that defined acculturation as language preference found that use of English related to use of oral contraceptives[13] and to fewer perceived barriers to medical care.[14] Further, a recent study[15] that examined several dimensions of acculturation (language, country of origin, contact with homeland, parental expectations of their children) found that language was the only dimension that predicted use of breast cancer screening among elderly Hispanic women; those who used English went for screening more recently. Taken together, these studies suggest that aspects of language may be especially important to our understanding of Hispanics' use of medical screening programs.

A second hypothesis attributes use of preventive health services to access to care. Access can be defined in terms of factors that influence the ease with which medical care can be obtained.[5] Such factors include the availability of health insurance, having a routine place of care and a regular provider, type of health care facility used and its proximity to residence.[5]

Compared to the general population, Hispanics are less likely to have health insurance coverage and less likely to have a routine place for obtaining health services.[5,18–21] This reflects their low income, low education, and employment in positions that do not provide insurance benefits.[4,5,22,23] Because of these barriers, they are more likely to rely on public health facilities, hospital outpatient clinics, and emergency rooms,[20,21,24,25] and are not likely to see a particular physician on an ongoing basis.[20] The "working poor" represent a segment of the Hispanic population that is especially vulnerable to access barriers in that they are neither able to afford the rising costs of medical services nor eligible for assistance through governmental programs.[21,26]

The present study provided a systematic examination of acculturation and access to care as factors influencing use of preventive health services by Hispanics. Unlike prior research on Hispanic health which has usually sampled only Mexican Americans, we examined three groups of Hispanics: Mexican Americans, Cuban Americans, and Puerto Ricans. Further, we measured use of an array of preventive services (e.g., routine physical, dental, and eye examinations, Pap smear, and breast examination).

Method

Sample

The data for our analyses were collected in the Hispanic Health and Nutrition Examination Survey (HHANES) conducted by the National Center for Health Statistics (NCHS) from 1982 through 1984. This survey sampled Mexican Americans from five southwestern states (California, Arizona, New Mexico, Texas, and Colorado), Cuban Americans from Dade County (Miami) Florida, and Puerto Ricans from

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NOTES: Author affiliations are listed elsewhere under CONTRIBUTORS.

the New York City area. For detailed information about the complex sampling design see the methodological paper[27] at the beginning of this supplement and NCHS plan and operation of HHANES.[28]

Selected sociodemographic information, as well as analytic sample sizes, are presented in Table 1. Analyses were restricted to eligible Hispanic sample persons between the ages of 20 and 74 who provided data on each variable used in the study. There were few cases of missing data. Of the total number of Mexican Americans, Cuban Americans, and Puerto Ricans (ages 20–74) sampled, 89 percent, 92 percent, and 93 percent, respectively, were included in the analytic samples.

Measures

Preventive Behavior—Utilization of preventive services was measured by responses to queries on seven service-related behaviors. The first item listed below was considered indicative of a person's tendency to *plan* his/her use of services, while the other six measures concerned recency of use of specific types of medical examinations.

- Individuals were asked whether or not they had made an appointment for their last visit to a health care setting.
- "About how long has it been since you had a *routine* physical examination, that is, not for a particular illness but for a general check-up?" (Responses were coded on a five-point scale ranging from "never" [1] to "less than one year ago" [5].)
- "About how long has it been since you last saw a dentist or dental hygienist for dental care?" (Responses were coded from "never" [1] to "six months ago or less" [6].)
- "When was the last time you had your vision tested?" ("Never" [1] to "six months ago or less" [6].)
- "About how long has it been since you last had your blood pressure taken?" ("Never" [1] to "six months ago or less" [4].)
- "About how long has it been since you had a Pap smear test for cancer?" ("Never" [1] to "less than one year ago" [8], females only.)
- "About how long has it been since you had your breasts examined by a doctor or other health professional?" ("Never" [1] to "less than one year ago" [8], females only.)

Preliminary analyses revealed that, for each sample, responses to these seven measures were positively correlated (r .11 to .74). Thus, we created a preventive index for each respondent separately for each sample. Each of the seven items was standardized to a mean of zero and a variance of one to give each equal weight. For the females, the seven standardized items were averaged to form the index; for the males, each of the five relevant standardized items was averaged to form the index. We should note that the HHANES did not examine use of health services over time; thus recency of use of the aforementioned services was the best indicator of regularity of use.

Acculturation– This construct was assessed with eight items pertaining to spoken and written language and to ethnic identification. The items represent a subset of the 20-item "Acculturation Rating Scale for Mexican Americans" developed by Cuellar, *et al.*[29]

1. What language do you speak?
2. What language do you prefer?
3. What language do you read better?
4. What language do you write better?

These four language measures, as well as the four ethnic measures below, were coded so that higher scores reflected a stronger English language/American orientation. Items 1–4 ranged from "Spanish only" [1], to the midpoint "Spanish

TABLE 1—Sample Characteristics

	Mexican Americans		Cuban Americans		Puerto Ricans	
Age range (years)	20–74		20–74		20–74	
Total sample (20–74)	3326		865		1220	
Analytic sample	2975		794		1138	
Men	1318		351		418	
Women	1657		443		720	
Marital Status	*M*	*F*	*M*	*F*	*M*	*F*
% Married– spouse in home	76.0	65.5	74.8	65.2	61.8	38.6
% Married—spouse not home	2.0	1.9	1.5	1.6	2.5	4.9
% Widowed	1.1	6.4	0.4	7.8	0.3	5.6
% Divorced	3.7	7.6	7.4	11.9	7.0	13.1
% Separated	1.7	6.5	1.5	3.0	4.3	13.8
% Never married	15.5	12.1	14.4	10.5	24.1	24.0
Age (years)						
% 20–30	41.6	39.4	23.2	20.8	34.5	35.8
% 31–40	28.0	26.0	17.2	22.6	26.6	27.5
% 41–50	13.1	14.0	22.7	22.3	18.1	19.1
% 51–74	17.3	20.6	36.9	34.3	20.8	17.6
Annual Family Income						
% Less than $10,000	23.3	33.9	20.7	29.5	29.5	45.0
% $10,000–$19,999	37.6	33.7	33.4	29.7	34.6	29.4
% $20,000–$29,999	21.2	18.0	21.9	17.5	17.2	10.3
% $30,000 and above	17.9	14.4	24.0	23.3	18.7	15.3
Education						
% Less than 7 years	28.1	30.9	22.5	22.6	16.7	16.8
% 7–11 years	26.7	28.6	24.7	22.2	36.1	35.9
% 12 years	23.7	24.2	20.7	26.6	28.2	30.1
% 13 or more	21.5	16.3	32.1	28.6	19.0	17.2

The percentages for marital status, age, income, and education were calculated with normalized sample weights (examination weights).

and English equally" [3], to "English only" [5]. The four remaining items measured ethnic identification:

5. What ethnic identification do you use?
6. What ethnic identification does/did your mother use?
7. What ethnic identification does/did your father use?
8. Where was the birthplace of yourself, your mother, your father?

For the Mexican American sample, the response codes for items 5–7 ranged from "Mexican" [1], to "Mexican American" [3], to "Anglo or other" [5]. For the Cuban Americans and Puerto Ricans, responses to items 5–7 had highly restricted variance; almost all of the Cuban Americans (and their parents) identified themselves as "Cuban"; the same situation occurred for the Puerto Ricans. Thus, these items could not be used meaningfully for these two groups. Item 8 was coded to establish generation. The scale ranged from "first generation (respondent born in country of origin)" [1] to "third or higher generation (respondent and both parents born in United States)" [4]. This item could not be used for the Cuban Americans and Puerto Ricans because of restricted variance; 97.3 percent of the Cuban Americans (ages 20–74) and 80 percent of the Puerto Ricans (ages 20–74) were first generation. Thus, for these two samples, only the four language items were used to assess acculturation. For the Mexican Americans, all eight items were used.

We constructed separate indices of acculturation for each respondent. Because we were interested in analyzing the effects of the different dimensions of acculturation and not just acculturation in general, we performed a factor analysis with the Mexican American data. It suggested that the items could be meaningfully grouped into one of three factors: (a) language preferred and spoken (items 1 and 2); (b) language written and read (items 3 and 4); and (c) ethnic identification (items 5–8). Each Mexican American respondent was assigned a score on each of these three dimensions by first standardizing the items that grouped together on the particular dimension and then averaging the items. Additionally, an overall score was created by averaging all eight items. For the Cuban Americans and Puerto Ricans separately, items 1 and 2 were standardized and then averaged to form a prefer/speak score, and items 3 and 4 were combined in similar fashion to form a read/write score. An overall score for the Cuban Americans and Puerto Ricans was calculated by averaging the four language items.

Access to health care—Five indicators of accessibility of health services were examined.

1. Health insurance, including private as well as public and federal/military assistance (e.g., Medicare, Medicaid, other public assistance, CHAMPUS*, other military programs). Any form of coverage was coded 1; no coverage was coded 0.
2. Whether or not the respondent had a routine place he or she used for health care (yes = 1, no = 0).
3. Type of health care facility used (private clinic, doctor's office, health maintenance organization, or prepaid group [4]; community clinics [3]; hospital outpatient clinic [2]; hospital emergency room [1]). The distribution of response across these categories was somewhat skewed (see Table 3); thus, for the regression analyses categories 3 and 4 were combined and assigned a code of 2, and categories 1 and 2 were combined and assigned a code of 1.
4. Whether or not the respondent usually sees a particular provider at the facility (yes = 1, no = 0).
5. Distance between residence and facility used (less than 10 minutes [3]; between 10 and 20 minutes [2]; over 20 minutes [1]).

An overall access score was constructed by standardizing and then averaging the five measures. In addition, for some analyses the individual measures of access were examined separately.

Data Analysis

All analyses were performed separately for each of the three samples, and separately for men and women applying normalized sample weights (examination weights).[27] We examined the weighted prevalence of preventive behavior and the weighted prevalence of access to services. Associations among variables were examined with weighted Pearson correlations. Our primary analyses of the effects of acculturation and access on recency of utilization were performed with multiple regression. Several regression models were tested, using as predictors the overall acculturation and overall access scores, and in different models using the individual components of these overall scores. Sociodemographic factors (age, education, income) were also included in the equation. Prior to conducting these tests, a statistical program (SESUDAAN)[30] was run to calculate the average design effect for each predictor variable. This average design effect was used to conservatively adjust the standard error of a regression coefficient generated by a weighted ordinary least-squares analysis. We should emphasize that, although certain variables were examined as "predictors," the HHANES is a cross-sectional design; thus, it does not provide evidence regarding cause and effect.

Results

Prevalence of Health Service Utilization

For each sample, recent use of the preventive health services was more prevalent among females than males (Table 2). Further, except for Pap smear and breast examination, recent use was higher among Cuban Americans and Puerto Ricans than among Mexican Americans. For example, less than 50 percent of the Mexican Americans reported that they had a routine physical examination within the last two years, and approximately 20 percent of the Mexican Americans reported that they never had a regular check-up from a physician. A similar pattern occurred for eye and dental examinations; recency of the female examinations was comparable among the samples.

Access to Health Services

The data suggest an association between access to care and utilization of health services. As seen in Table 3, a greater percentage of women than men reported that they had a routine place for health care and that they usually saw a particular provider at the facility. This gender difference in access parallels the gender differences in utilization described above. Second, fewer Mexican Americans (compared to the other two groups) had some form of health insurance coverage (Table 3) which may partially explain their low rate of screening. Finally, consistent with prior research,[31] a substantially larger percentage of Puerto Ricans compared to

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*Civilian Health and Medical Program of the Uniformed Services.

TABLE 2—Weighted Prevalence of Utilization of Services

	Mexican Americans		Cuban Americans		Puerto Ricans	
	Male	Female	Male	Female	Male	Female
Last routine physical						
Less than 2 years ago	33.1	45.4	42.8	59.5	47.3	61.6
Over 2 years ago	46.4	23.8	45.2	31.6	45.3	32.7
Never	20.5	20.8	12.0	8.9	7.4	5.7
Made appointment for visit						
Yes	59.7	71.7	62.8	75.3	45.4	59.4
Last blood pressure taken						
Less than 1 year ago	57.7	78.9	77.7	81.3	72.8	85.0
Over 1 year ago	36.4	19.9	20.3	18.1	25.4	14.6
Never	5.9	1.2	2.0	0.6	1.8	0.4
Last eye examination						
Less than 2 years ago	51.3	51.2	63.5	63.6	62.7	66.1
Over 2 years ago	36.3	38.2	31.8	31.0	31.0	29.7
Never	12.3	10.6	4.7	5.4	6.3	4.2
Last dental check-up						
Less than 2 years ago	49.7	60.0	62.7	72.1	60.8	74.8
Over 2 years ago	39.0	34.0	36.2	27.7	37.0	24.6
Never	11.3	6.0	1.1	0.2	2.2	0.6
Last PAP smear						
Less than 2 years ago		75.3		68.2		79.6
2-5 years ago		9.3		14.0		7.1
Over 5 years or never		15.4		17.8		13.3
Last breast examination						
Less than 2 years ago		77.9		73.2		83.0
2-5 years ago		8.0		11.1		7.0
Over 5 years or never		14.1		15.7		10.0

Entries are not age-adjusted.

Mexican Americans and Cuban Americans used hospital outpatient clinics and hospital emergency rooms as their usual source of health care (Table 3).

Access was correlated with sociodemographic factors. For Mexican American and Cuban American men and women, the overall access score correlated positively with education, annual income, and age (r .10 to .25). The correlation between health insurance coverage and income was especially strong (r .24 to .37). A similar pattern of results was found for Puerto Ricans, but the correlations were not as strong.

Regression Analyses

Overall Effects—The first set of multiple regression analyses used the overall acculturation score, the overall access score, and sociodemographic factors (education, income, age) as predictors of the preventive service use index. The predictors were entered into the equation simultaneously; thus, the effect of each variable was assessed after controlling for the effects of each of the other variables in the equation. As seen in Table 4, the R^2 from each analysis indicated that the array of predictors accounted for more

TABLE 3—Access Related Variables (weighted prevalence)

	Mexican Americans		Cuban Americans		Puerto Ricans	
	Male	Female	Male	Female	Male	Female
Health insurance						
Yes	65.9	66.1	72.1	75.3	70.0	82.7
No	34.1	33.9	27.9	24.7	30.0	17.3
Routine place for health care						
Yes	64.9	82.2	60.7	77.9	74.6	83.4
No	35.1	17.8	29.3	22.1	25.4	16.6
Type of facility used						
Doctor's office, Private clinic, HMO/prepaid group	69.2	70.7	71.8	76.9	46.0	55.1
Community clinic	16.7	19.5	15.5	14.5	18.9	18.9
Hospital outpatient clinic	8.9	8.1	7.8	7.5	25.5	22.1
Hospital ER	5.2	1.7	4.9	1.1	9.6	3.9
Particular provider						
Yes	58.0	69.3	59.6	72.3	47.6	64.3
No	42.0	30.7	40.4	27.7	52.4	35.7
Distance from health care facility						
< 10 min	70.6	78.2	69.5	74.6	71.1	73.8
10-20 min	18.2	13.0	19.5	16.5	17.0	17.7
over 20 min	11.2	8.8	11.0	8.9	11.9	8.5

Entries are not age-adjusted.

TABLE 4—Regression Analyses of Acculturation, Access, and Sociodemographics Predicting Preventive Behavior

	Mexican Americans				Cuban Americans				Puerto Ricans			
	Males		Females		Males		Females		Males		Females	
	b	SE	b	SE	b	SE	b	SE	b	SE	b	SE
Acculturation[a]	.100	(.033)	.042	(.023)	.121	(.042)	.059	(.043)	.082	(.086)	−.033	(.027)
Access to care[b]	.182	(.032)	.194	(.021)	.323	(.040)	.207	(.039)	.352	(.063)	.289	(.047)
Education	.032	(.005)	.026	(.005)	.018	(.009)	.013	(.008)	.034	(.010)	.018	(.008)
Annual income	.007	(.003)	.008	(.003)	.001	(.004)	.001	(.006)	−.002	(.004)	.005	(.003)
Age	.009	(.001)	.002	(.016)	.006	(.002)	.004	(.002)	.014	(.003)	.004	(.002)
	R^2 = .212		R^2 = .161		R^2 = .262		R^2 = .104		R^2 = .239		R^2 = .142	

Entries are weighted regression coefficients and adjusted standard errors in parentheses.
a) Overall acculturation score.
b) Overall access score.

variance in the men's preventive behavior than the women's. With respect to individual effects, access to health services was a uniformly strong predictor in each sample. Acculturation, however, was an important predictor only for Mexican American and Cuban American men. The regression coefficients for the sociodemographic variables were small, but the pattern indicated that increased education, increased income, and older age were associated with greater recency in utilization of the health services.

Specific Effects of Acculturation—Additional regression analyses were performed to examine the effects of the individual components of the overall acculturation score (spoken language, written language, and ethnic identification). Because these three components were moderately to highly correlated (*r* .53 to .81), analyzing the three together in a simultaneous model could generate misleading results due to the multicollinearity among the variables. Thus, in addition to testing a simultaneous model, we tested three additional models in which the acculturation measures were analyzed separately. The effects of ethnic identification could be examined only for the Mexican Americans. All of the analyses controlled for age, income, and education.

As seen in Table 5, for Mexican American men and women, Model 1 (all three acculturation variables analyzed together) indicated that only spoken language was significant. That is, preferring to speak English was associated with greater recency in utilization of the services. Models 2, 3, and 4 (which analyzed the three acculturation variables separately) yielded results consistent with the first model. Although spoken language (Model 2) and written language (Model 3) each were predictors, spoken language was the stronger predictor. Ethnic identification was not a predictor in either the simultaneous or individual model (Model 4).

The findings for Cuban Americans and Puerto Ricans were not as clear. For Cuban American men, written language produced stronger effects than spoken language, and neither language dimension produced appreciable effects for Cuban American women or for Puerto Rican men and women.

Specific Effects of Access—Not surprisingly, some of the access variables were correlated. For each sample, having a particular provider was correlated with using HMOs or private clinics (*r* .28 to .48) and with having a routine place for health care (*r* .34 to .60). The correlations among the other access variables were much smaller.

A series of regression analyses were performed that tested simultaneous as well as individual models for the specific access variables. The simultaneous model (Model 1) indicated that, for each sample, having a routine place for health care was generally the strongest predictor of recency in utilization (see Table 6). Comparison of the individual models also reveals that having a routine place for care produced the strongest effects (Model 3). The findings for health insurance were strong in both the simultaneous and the individual model (Model 2), and having a particular provider was a predictor in the individual model (Model 5). The results for the other measures of access were less consistent. Type of health care facility used most often was a predictor only for the men of each sample and only when this variable was the sole access variable in the equation (Model 4). Finally, distance from facility was predictive of utilization only for Puerto Ricans (Model 6); as distance decreased, recency of screening increased.

Discussion

Our analyses unveiled interesting differences in preventive service-related behaviors among the three Hispanic groups. Mexican Americans used such services less recently than the other two groups, except in the case of female examinations. Previous research has also shown that Mexican Americans are less likely than non-Hispanic Whites and Blacks to visit a physician or a dentist, or to be admitted to a hospital.[32] In part, this can be explained by the fact that Mexican Americans are less likely than others to have health insurance as demonstrated in the present and other studies.[33]

There are several reasons for the higher prevalence of health insurance coverage among Cuban Americans and Puerto Ricans. Cuban Americans compared to Mexican Americans are more likely to have private insurance coverage as a result of higher incomes and/or greater availability of health insurance as an employee benefit.[33] Puerto Ricans, however, are more likely than Mexican Americans to be covered by Medicaid; they have higher proportions of female-headed households eligible for Medicaid under Aid to Families with Dependent Children and are more likely to reside in states that provide greater optional coverage under Medicaid.[26,33] The higher prevalence of Medicaid coverage among Puerto Ricans may also explain their greater use of public facilities.

At the outset, we offered two hypotheses to explain the use of preventive health services by Mexican Americans, Cuban Americans, and Puerto Ricans; one emphasized the importance of acculturation and the other emphasized access

TABLE 5—Components of Acculturation Predicting Preventive Behavior

	Males				Females			
Models:	1	2	3	4	1	2	3	4
Mexican Americans								
Spoken language	.116	.139	—	—	.105	.087	—	—
	(.028)	(.022)			(.023)	(.016)		
Written language	.042	—	.104	—	−.027	—	.047	—
	(.032)		(.037)		(.024)		(.016)	
Ethnic ID	−.010	—	—	.059	.004	—	—	.039
	(.026)			(.037)	(.047)			(.042)
Education	.031	.033	.036	.044	.023	.023	.029	.032
	(.005)	(.005)	(.005)	(.005)	(.005)	(.005)	(.005)	(.004)
Annual income	.010	.009	.011	.012	.012	.012	.012	.013
	(.003)	(.003)	(.003)	(.003)	(.003)	(.003)	(.003)	(.003)
Age	.012	.012	.012	.012	.004	.004	.004	.004
	(.002)	(.002)	(.002)	(.002)	(.001)	(.001)	(.001)	(.002)
Cuban Americans								
Spoken language	−.021	.081	—		.018	.058	—	
	(.051)	(.037)			(.043)	(.033)		
Written language	.152	—	.137		.060	—	.072	
	(.052)		(.038)		(.087)		(.078)	
Education	.024	.031	.023		.012	.015	.013	
	(.009)	(.009)	(.009)		(.008)	(.008)	(.008)	
Annual income	.007	.009	.007		.004	.004	.004	
	(.007)	(.007)	(.006)		(.007)	(.008)	(.007)	
Age	.012	.010	.012		.006	.005	.006	
	(.004)	(.003)	(.004)		(.003)	(.002)	(.003)	
Puerto Ricans								
Spoken language	.063	.046	—		.045	−.011	—	
	(.079)	(.092)			(.061)	(.025)		
Written language	.011	—	.053		−.076	—	−.042	
	(.100)		(.081)		(.079)		(.025)	
Education	.043	.038	.044		.022	.018	.024	
	(.011)	(.010)	(.011)		(.008)	(.007)	(.008)	
Annual income	.003	.003	.004		.007	.006	.008	
	(.005)	(.005)	(.005)		(.003)	(.003)	(.003)	
Age	.019	.018	.019		.007	.007	.007	
	(.003)	(.003)	(.003)		(.002)	(.002)	(.002)	

Entries are weighted regression coefficients and adjusted standard errors in parentheses. Variables not entered into a specific regression model are indicated by a dash (—). Ethnic identification was not examined for Cuban Americans or Puerto Ricans.

to services. Our findings clearly demonstrated that access was strongly related to utilization. In particular, having a routine place for care and having health insurance coverage produced the strongest effects. With respect to acculturation, language but not ethnic identification was the important predictor of utilization for the Mexican Americans. Language was related to utilization for the Cuban American men but not for the Cuban American women or Puerto Ricans. We are uncertain why the language effect was not found among these latter Hispanic subgroups. In the case of the Puerto Ricans, the public health facilities they tend to use (because of their high rate of Medicaid coverage) may have more Spanish-speaking providers than private clinics and hospitals so that use of the English language may not be essential to access.

Despite the differential effect of language among the three Hispanic groups, language was clearly associated with recency of screening. A prior study obtained similar results.[15] Of four dimensions of acculturation (language, country of origin, contact with homeland, parental expectations of children) only language was significantly associated with use of screening examinations (e.g., physician breast examination, breast self-examination). Further, as discussed earlier, several studies have found that service utilization rates were higher among non-Hispanic Whites than Hispanics after controlling for sociodemographic factors and health insurance coverage.[3–5,11] Although this difference has been attributed to cultural factors in the broad sense, the present data suggest that the results were probably due to specific language differences between the groups.

What interpretation should be placed on this language effect? Traditionally, the effect of language has been viewed as an indicator of the importance of cultural factors in health service utilization. If cultural factors were important, one would expect to find strong effects of language as well as significant effects of the psychological and social dimensions of acculturation. These latter dimensions, however, have not been shown to be associated with screening behavior. This causes us to question the usefulness of a broad cultural interpretation of the language effect. Our preferred view is that language represents an access factor. Although language preference can be viewed as a measure of cultural integration, it also can be seen as an indicator of functional integration.[34] From this latter perspective, ability to speak English increases the extent to which Hispanics can effectively attain institutional access. This view is consistent with several prior papers that have discussed language as a potential barrier to health service delivery.[7–9] In sum, regardless of one's level of acculturation on psychological or social dimensions, variation in language preference seems to be a critical determinant of utilization of health services, and is best viewed in terms of accessibility.

Recommendations

The present results support several ideas discussed in the literature for increasing the use of health services by

TABLE 6—Components of Access to Care Predicting Preventive Behavior

	Males						Females					
Models:	1	2	3	4	5	6	1	2	3	4	5	6
Mexican Americans												
Health insurance	.068 (.024)	.204 (.049)	—	—	—	—	.116 (.036)	.160 (.034)	—	—	—	—
Routine place for health care	.085 (.047)	—	.290 (.108)	—	—	—	.298 (.079)	—	.357 (.034)	—	—	—
Type of facility used	.166 (.100)	—	—	.216 (.048)	—	—	−.022 (.042)	—	—	.050 (.042)	—	—
Particular provider	.010 (.024)	—	—	—	.210 (.078)	—	.007 (.036)	—	—	—	.164 (.045)	—
Distance to facility	−.023 (.019)	—	—	—	—	.051 (.043)	.024 (.042)	—	—	—	—	.143 (.020)
Education	.038 (.004)	.046 (.004)	.044 (.004)	.044 (.004)	.045 (.004)	.049 (.004)	.031 (.004)	.033 (.003)	.033 (.004)	.034 (.004)	.033 (.004)	.034 (.003)
Annual income	.007 (.003)	.007 (.003)	.010 (.002)	.011 (.002)	.010 (.003)	.012 (.003)	.008 (.002)	.010 (.003)	.010 (.003)	.013 (.002)	.011 (.003)	.012 (.003)
Age	.008 (.002)	.011 (.002)	.010 (.001)	.011 (.002)	.011 (.002)	.012 (.002)	.002 (.002)	.003 (.002)	.003 (.002)	.004 (.001)	.003 (.002)	.004 (.002)
Cuban Americans												
Health insurance	.217 (.067)	.322 (.091)	—	—	—	—	.133 (.152)	.172 (.125)	—	—	—	—
Routine place for health care	.419 (.083)	—	.537 (.058)	—	—	—	.263 (.085)	—	.301 (.057)	—	—	—
Type of facility used	.121 (.088)	—	—	.252 (.089)	—	—	.013 (.150)	—	—	.116 (.148)	—	—
Particular provider	.047 (.102)	—	—	—	.358 (.120)	—	.142 (.063)	—	—	—	.258 (.053)	—
Distance to facility	−.052 (.070)	—	—	—	—	.129 (.083)	−.089 (.047)	—	—	—	—	.061 (.043)
Education	.029 (.008)	.034 (.008)	.033 (.008)	.035 (.008)	.036 (.008)	.038 (.008)	.018 (.007)	.020 (.007)	.022 (.007)	.020 (.007)	.020 (.007)	.021 (.007)
Annual income	.001 (.004)	.005 (.006)	.006 (.004)	.008 (.005)	.006 (.007)	.009 (.007)	.001 (.006)	.002 (.008)	.005 (.008)	.005 (.007)	.002 (.006)	.006 (.008)
Age	.001 (.002)	.006 (.002)	.003 (.002)	.007 (.002)	.006 (.002)	.008 (.003)	.001 (.002)	.003 (.002)	.004 (.002)	.004 (.002)	.003 (.002)	.005 (.002)

Hispanics. Because our data indicate that Mexican Americans are the least likely of the three Hispanic groups to be recent users of preventive services, the following recommendations are particularly relevant to them.

- First, special outreach efforts need to be conducted with Hispanics who are monolingual Spanish or prefer to speak Spanish regarding the importance of medical screening for prevention of illness. Programs directed toward medical providers regarding language issues, increasing bilingual staff, and programs dealing with patient-provider interaction would also be beneficial.
- Second, the present data showed that those with a regular source of care and a particular provider (e.g., family doctor, private clinic, prepaid group) were most likely to use preventive services. Because large portions of Hispanics do not have a regular provider, it would seem important either to: (a) incorporate screening programs into the settings that are used by them, or (b) establish effective referral mechanisms to direct individuals to screening centers.
- Finally, lack of health insurance is a barrier to the use of preventive services, particularly among Mexican Americans. This suggests the need to: 1) increase the availability of low-cost or no-cost medical screening programs in low income Hispanic communities; 2) increase the availability of health insurance coverage as an employee benefit; 3) broaden the eligibility criteria for Medicaid (which in many states is re-

TABLE 6—Continued

	Males						Females					
Models:	1	2	3	4	5	6	1	2	3	4	5	6
Puerto Ricans												
Health insurance	.295 (.138)	.346 (.133)	—	—	—	—	.174 (.109)	.251 (.079)	—	—	—	—
Routine place for health care	.496 (.086)	—	.520 (.114)	—	—	—	.306 (.065)	—	.387 (.050)	—	—	—
Type of facility used	.165 (.089)	—	—	.163 (.076)	—	—	.053 (.115)	—	—	.127 (.077)	—	—
Particular provider	.023 (.069)	—	—	—	.257 (.104)	—	.112 (.047)	—	—	—	.243 (.051)	—
Distance to facility	−.044 (.050)	—	—	—	—	.132 (.055)	.002 (.053)	—	—	—	—	.134 (.063)
Education	.036 (.009)	.041 (.010)	.040 (.009)	.036 (.010)	.041 (.010)	.042 (.010)	.015 (.006)	.017 (.007)	.015 (.006)	.017 (.007)	.015 (.007)	.018 (.007)
Annual income	−.001 (.004)	.001 (.005)	.005 (.004)	.003 (.005)	.003 (.005)	.004 (.004)	.005 (.003)	.006 (.003)	.007 (.003)	.005 (.003)	.004 (.003)	.005 (.003)
Age	.011 (.002)	.015 (.003)	.014 (.002)	.016 (.003)	.016 (.002)	.017 (.003)	.005 (.002)	.007 (.002)	.005 (.002)	.007 (.002)	.006 (.002)	.007 (.002)

Entries are weighted regression coefficients and adjusted standard errors in parentheses. Variables not entered into a specific regression model are indicated by a dash (—).

stricted to single-parent families); and 4) set satisfactory remuneration rates for health care facilities that give uncompensated care to the working poor.[26]

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III. Health Risk Behaviors of Hispanics in the United States: Findings from HHANES, 1982–84

Gary Marks, PhD, Melinda Garcia, PhD, and Julia M. Solis, PhD

Abstract: With data from the Hispanic Health and Nutrition Examination Survey (HHANES), we examined several health risk behaviors (cigarette smoking, alcohol use, dietary practices, and recency of health screening) of Mexican American, Cuban American, and Puerto Rican adults (ages 20–74). For each sample, a greater percentage of men than women smoked cigarettes and used alcohol. Heavy smoking (20+ cigarettes per day) was most prevalent for Cuban American males, and heavy drinking (1.00+ oz ethanol per day) was most prevalent for Mexican American and Puerto Rican men. Acculturation correlated positively with alcohol use (particularly for females) and negatively with dietary balance (for Mexican American men and women). The Puerto Ricans' diet was less balanced than that of the other two groups. For each sample, more men than women had not had a routine physical or dental examination within the past five years; the recency of screening was lowest for Mexican American men. Screening (including Pap smear for the women) was lower for those who smoked cigarettes and for those with poor dietary practices, indicating that many Hispanics at special risk of disease underutilize preventive health services, increasing the likelihood of diagnosis at a later stage of illness. [*Am J Public Health* 1990; 80(Suppl):20–26.]

Introduction

Compared to non-Hispanic Whites in the United States, Mexican Americans underutilize preventive health services such as routine physical check-ups, dental and eye examinations, and prenatal care.[1–5] When they present with an illness, the illness frequently is at a later stage.[6,7] Consequently, Mexican Americans are at greater risk of morbidity and mortality than the general United States population.[8]

The screening behavior of persons of Mexican heritage has been shown to be affected by: *access to care*, such as having a regular family physician and health insurance coverage;[2,3,5,9] *cultural variables* such as adaptation to the practices of American society;[5,10–12] and *sociodemographic factors* such as income, education, age, and gender.[13,14] The extent to which utilization of preventive health services is influenced by attitudinal and behavioral factors, however, has received much less attention.

Attitudes about health are reflected in many behaviors. For example, those who smoke cigarettes heavily,[15–17] drink alcohol excessively,[18,19] or maintain an unbalanced and unvaried diet (e.g., high in fried meats and low in fruits, vegetables, and fiber[17,20,21]) place themselves at increased risk of specific diseases, such as cancer of the lungs, pancreas, colorectal area, and prostate,[17,20–25] and hypertension and coronary heart disease.[26,27] Ideally, individuals who engage in these "health risk-behaviors" should use screening services regularly in order to monitor their health status. We suspect, however, that (after controlling for income, education, age, and acculturation) they use those services *less* frequently than individuals who do not engage in such behaviors, based on the assumption that such behaviors are components of a constellation that emanates from a general attitude toward health promotion and disease prevention.[28] This cluster of behaviors is likely to contribute to the presence of illness and to diagnosis at a later stage.

We provide data on the prevalence of these health risk behaviors among Mexican American, Cuban American, and Puerto Rican adults, and examine the extent to which recency of screening (physical and dental examination, Pap smear) is correlated with them.

Methods

Sample

The study used data from the Hispanic Health and Nutrition Examination Survey (HHANES), conducted by the National Center for Health Statistics (NCHS) from 1982–1984. This survey sampled Mexican Americans from five southwestern states, Cuban Americans from Dade County (Miami) Florida, and Puerto Ricans from the New York City area. For detailed information about the complex sampling design see the methodological paper by Delgado, *et al*,[29] at the beginning of this supplement and the NCHS plan and operation of HHANES.[30]

Our analyses focused on adults (20–74 years of age) from each of these three samples. Table 1 presents the total sample sizes in this age range, the analytic sample sizes, and selected sample characteristics. Analyses were restricted to eligible Hispanic sample persons who provided data on each variable used in this study. Of the total number of Mexican Americans, Cuban Americans, and Puerto Ricans (20–74) sampled, 85 percent, 86 percent, and 87 percent, respectively, were included in the analytic samples.

Measures

Health Service Utilization—Recency of physical examination was measured by asking "How long has it been since you had a *routine* physical examination, that is, not for a particular illness, but for a general check up?" Responses were coded on a 5-point scale ranging from "never" [1] to "less than one year ago" [5]. Recency of dental examination was assessed by asking how long it was since they saw a dentist or dental hygienist for dental care (coded from "never" [1] to "six months ago or less" [6]). Recency of Pap smear was examined for the women ("never" [1] to "less than one year ago" [7]).

Cigarette Smoking—This behavior was measured two ways: (a) a dichotomous measure reflected whether or not the respondent was a current cigarette smoker (current smoker = 1; does not currently smoke cigarettes, cigars, or pipes = 0; the few persons who smoked cigars or pipes only were not included in the analyses); or (b) among current cigarette smokers, each was assigned a score indicating the number of cigarettes smoked per day. These two measures allowed us to examine whether the effects of smoking are due to mere

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NOTE: Author affiliations are listed elsewhere under CONTRIBUTORS.

TABLE 1—Sample Characteristics

	Mexican Americans		Cuban Americans		Puerto Ricans	
Age range (years)	20–74		20–74		20–74	
Total sample (20–74)	3326		865		1220	
Analytic sample size	2824		745		1066	
Men	1249		330		389	
Women	1575		415		677	
	M	F	M	F	M	F
Marital Status						
% Married—spouse in home	76.5	65.8	74.6	65.6	60.5	39.0
% Married—spouse not home	2.0	1.9	1.3	1.6	2.8	5.1
% Widowed	1.1	6.3	0.5	7.8	0.3	5.8
% Divorced	3.7	7.5	7.8	12.0	7.5	13.4
% Separated	1.8	6.4	1.6	3.1	4.4	14.3
% Never married	14.9	12.1	14.2	10.0	24.5	22.4
Age (years)						
% 20–30	41.4	39.5	23.6	20.1	35.0	34.6
% 31–40	28.0	25.9	17.6	22.9	27.6	27.9
% 41–50	13.3	14.1	23.5	22.9	16.3	19.6
% 51–74	17.3	20.5	35.3	34.1	21.1	17.9
Annual Family Income						
% Less than $10,000	23.1	33.6	20.6	29.0	29.2	44.6
% $10,000–$19,999	37.9	33.7	33.4	30.3	34.8	29.0
% $20,000–$29,999	21.1	18.1	22.6	17.7	17.4	10.9
% $30,000 and above	17.9	14.6	23.4	23.0	18.6	15.5
Education						
% Less than 7 years	27.9	30.9	22.2	22.6	16.3	16.4
% 7–11 years	26.6	28.5	24.9	21.9	36.1	35.9
% 12 years	24.1	24.2	21.3	27.2	28.4	30.1
% 13 or more	21.4	16.5	31.6	28.3	19.2	17.6

The percentages for marital status, age, income, and education were calculated with normalized sample weights (examination weights).

smoking status (smoker vs nonsmoker) or to magnitude of smoking among current smokers.

Alcohol Use—Alcohol consumption was grouped into four volume levels: (a) little or no current use (fewer than 12 drinks of beer, wine, or liquor in one's entire life or in any one year, or no current use); (b) light regular/occasional drinkers (light regular drinkers consumed an average of 0.01 to 0.21 oz of ethanol per day, occasional drinkers drank in the light ethanol range or above but did so irregularly); (c) moderate regular drinkers (an average of 0.22 to 0.99 oz of ethanol per day); and (d) heavy regular drinkers (an average of 1.00 or more oz of ethanol per day). The ethanol scale was constructed by NCHS using a conversion factor in which one oz of beer contains 0.04 oz of ethanol, one oz of wine contains 0.15 oz of ethanol, and one oz of hard liquor contains 0.45 oz of ethanol.

Diet Index—Diet was assessed with a food item and consumption frequency checklist similar to that used by others.[17,20,22,23] Respondents reported on all food items consumed in the past three months. They indicated whether they had eaten the food item daily (coded as a score of 3), weekly (2), less than once a week (1), or never (0). We constructed an overall diet index that consisted of two components. One component was the extent to which an individual's diet was balanced. Specific food items were grouped into dairy products, meats (including poultry and fish), fruits, vegetables, and grains/cereals. For each person, scores were averaged separately for each food group, and then the five subscale scores were averaged. Higher overall scores indicated a more balanced diet (i.e., frequent consumption of foods in each group). The second component consisted of a "junk food" factor, i.e., the extent to which an individual consumed items such as candy, sodas (excluding diet soda), cake, cookies, chips, sugar, etc. The overall diet index was formed by splitting the distribution of scores for dietary balance and the distribution of scores for "junk food" at their respective medians. Those high on dietary balance and low on "junk food" intake were assigned a score of 3; those high on each dimension received a score of 2; those low on each dimension received a score of 1; and those low on balance and high on "junk food" were assigned a score of 0. This index was used in the correlations and regression analyses. For descriptive purposes, we calculated the number of the five food groups from which respondents had eaten one or more food items daily, and the percentage who had eaten one or more "junk food" items daily.

Acculturation—This variable was measured with eight items pertaining to spoken and written language and to ethnic identification. The items represent a subset of the 20-item "Acculturation Rating Scale for Mexican Americans" developed by Cuellar, *et al.*[31]

1. What language do you speak?
2. What language do you prefer?
3. What language do you read better?
4. What language do you write better?

These four measures (as well as the four below) were coded so that higher scores reflected a stronger English language/American orientation. Items 1–4 ranged from "Spanish only" (1), to the midpoint "Spanish and English equally" (3), to "English only" (5). The four remaining items measured ethnic identification:

5. What ethnic identification do you use?
6. What ethnic identification does/did your mother use?
7. What ethnic identification does/did your father use?
8. Where was the birthplace of yourself, your mother, your father?

For the Mexican American sample, the response codes for items 5–7 ranged from "Mexican" (1), to "Mexican American" (3), to "Anglo or other" (5). For the Cuban Americans and Puerto Ricans, responses to items 5–7 had highly restricted variance; almost all of the Cuban Americans (and their parents) identified themselves as "Cuban"; the same situation occurred for the Puerto Ricans. Thus, these items could not be used meaningfully for these two groups. Item 8 was coded to establish generation. The scale ranged from "first generation (respondent born in country of origin)" (1) to "third or higher generation (respondent and both parents born in US)" (4). This item could not be used for the Cuban Americans and Puerto Ricans again because of restricted variance; 97.5 percent of the Cuban Americans (age 20–74) and 80.5 percent of the Puerto Ricans (age 20–74) were first generation. Thus, for these two samples, only the four language items were used to assess acculturation. These four items were standardized to a mean of zero and variance of one (to give each item equal weight) and then averaged to create a composite score for the Cuban Americans and Puerto Ricans. For the Mexican Americans, all eight items were standardized and then averaged to create the acculturation score. Demographic information on annual income, education, and age was also included in the analyses.

Data Analysis

Analyses were performed separately for each of the three samples, and separately for men and women applying normalized sample weights (examination weights).[29] We analyzed the weighted prevalence of each health risk behavior as well as the weighted Pearson correlation of these behaviors with acculturation, education, income, and age. Our primary hypothesis was examined with multiple regression analysis. The regression model for each dependent measure (i.e., utilization variable) included as predictors the three health risk behaviors, acculturation, and the three sociodemographic variables. The predictors were entered into the equation simultaneously. Prior to conducting these tests, a statistical program (SESUDAAN)[32] was run to calculate the average design effect for each predictor variable. This average design effect was used to conservatively adjust the standard error of a regression coefficient generated by a weighted ordinary least-squares analysis. We should emphasize that although certain variables were examined as "predictors," the HHANES is a cross-sectional design; thus, it does not provide evidence regarding cause and effect.

Results

Prevalence of Health Risk Behaviors

For each sample, behaviors that increase one's risk of illness and disease were more prevalent for men than women. A greater percentage of the men smoked cigarettes and drank alcohol heavily (Table 2). Heavy cigarette smoking was most prevalent for Cuban American men, and heavy drinking most prevalent for Mexican American and Puerto Rican males. Dietary balance was somewhat lower for men than women, and lower for the Puerto Ricans compared to the other two groups. Further, for each sample, more men than women had not had a physical or dental examination within the last five years. These gender differences occurred for those above, as well as for those below, the median on the acculturation index. The utilization of the health services was especially low for the Mexican American men. Table 2 also indicates that use of the preventive services was lower for the Mexican Americans than the Cuban Americans and Puerto Ricans. One explanation for this difference is that fewer Mexican Americans had some form of health insurance coverage.[5]

Correlation of Health Risk Behaviors with Acculturation and Sociodemographics

With only a few exceptions, the pattern of the correlations was similar for men and women, and for each sample (Table 3). Because of the relatively large sample sizes, many of the coefficients were statistically significant even though their absolute magnitudes were low. There were, however, several meaningful effects. The strongest associations involved acculturation, which correlated with increased alcohol consumption (particularly for the women of each sample), and with lower diet scores for the Mexican American men and women. As age increased, alcohol consumption decreased; the correlations between age and cigarette smoking and between age and diet were small. Those with more

TABLE 2—Weighted Prevalence of Health Risk Behaviors

	Mexican Americans		Cuban Americans		Puerto Ricans	
	Males	Females	Males	Females	Males	Females
Percent current cigarette smoker	43.9	24.2	45.1	23.9	41.9	33.5
Percent smoking 20+ cigarettes per day	13.7	4.5	27.3	11.9	22.0	12.2
Percent current alcohol user	77.4	34.7	68.9	22.8	70.8	32.6
Percent drinking 1.00+ oz ethanol per day	17.1	1.2*	9.2	0.4*	17.1	2.5*
Dietary balance[a]						
0–1	22.4	16.9	14.8	13.0	29.0	30.3
2–3	50.9	51.0	57.3	52.1	50.1	49.5
4–5	26.7	32.1	27.9	34.9	20.9	20.2
"Junk food" intake[b]	67.4	61.0	75.3	75.7	79.1	73.9
Percent having last physical examination over 5 years ago or never	46.1	37.5	33.0	22.7	31.8	19.4
Percent having last dental examination over 5 years ago or never	29.9	21.7	16.5	9.6	19.1	9.5
Percent having last Pap smear over 5 years ago or never		15.4		17.8		13.3

Entries are not age-adjusted.

*An asterisk next to an entry indicates that the statistic should be viewed cautiously because it does not meet the minimum reliability standard due to small cell size.

[a] The scale for dietary balance indicates the number of healthy food groups from which the person had eaten one or more food items daily.

[b] The entries for "junk food" indicate the percentage who had eaten one or more "junk food" items daily.

TABLE 3—Correlation of Health Risk Behaviors with Acculturation and Sociodemographics

	Age		Education		Income		Acculturation	
	M	F	M	F	M	F	M	F
Mexican Americans								
Smoking	.01	.01	−.12	−.01	−.17	−.07	−.06	.09
Alcohol use	−.11	−.13	.12	.29	.10	.10	.11	.30
Diet index	.04	.08	.06	.03	.06	.05	−.12	−.12
Cuban Americans								
Smoking	−.10	−.11	−.04	.03	.04	−.05	−.15	−.01
Alcohol use	−.18	−.29	.10	.30	.13	.21	.13	.36
Diet index	−.01	.10	.08	.12	.05	.06	.04	.03
Puerto Ricans								
Smoking	−.06	−.13	−.13	.01	−.18	−.06	−.07	.17
Alcohol use	−.13	−.17	.05	.25	−.01	.12	.04	.26
Diet index	.03	.16	.07	.15	.11	.13	.06	−.03

Entries are weighted Pearson correlations. Cigarette smoking was a dichotomous variable with current smoker = 1 and nonsmoker = 0.

education and higher income (especially the women) consumed a greater volume of alcohol.

Correlates of Screening

Smoking—In separate regression analyses the two cigarette smoking measures were examined as predictors of recency of screening. Only the dichotomous measure was significant, which suggests that the key factor associated with screening activity was whether or not one smokes, not necessarily the amount one smokes. As seen in Table 4, among the Mexican Americans and Cuban Americans, smokers were less recent users of the health services than were the nonsmokers. These effects are seen more concretely in Table 5, which presents the weighted prevalence of recent utilization by smoking status (not controlling for other factors). The findings were strongest for the Mexican American and Cuban American men; for these two groups, the results were consistent across levels of acculturation. The relation between smoking and screening for the Puerto Rican men and women was generally in the predicted direction, although very weak.

Diet—Dietary practices were associated with screening, especially for the Mexican American women. Among this subgroup, those with poorer diets (i.e., low in balance and high in "junk food" intake) were less recent users of the health services (Tables 4 and 5). This association was not qualified by the degree of acculturation of the Mexican American women.

An interesting pattern emerged for the diet and smoking variables for the Mexican American sample. Whereas the

TABLE 4—Regression Analyses of Health Risk Behaviors, Acculturation, and Sociodemographics Predicting Recency of Screening

	Males				Females					
	Physical Examination		Dental Examination		Physical Examination		Dental Examination		Pap Smear	
	b	SE	b	SE	b	SE	b	SE	b	SE
Mexican Americans										
Smoking	−.139	.085	−.354	.103	−.085	.089	−.086	.092	.013	.107
Diet index	.101	.046	.043	.055	.130	.039	.169	.040	.110	.047
Alcohol	−.079	.039	.086	.047	−.050	.090	.119	.091	.108	.106
Acculturation	.244	.084	.117	.103	.119	.098	.011	.101	.166	.118
Education	.050	.020	.061	.025	.041	.023	.085	.023	.038	.027
Income	.015	.003	.031	.007	.013	.006	.029	.007	.025	.007
Age	.014	.004	.014	.004	.012	.004	.008	.004	−.011	.004
Cuban Americans										
Smoking	−.275	.158	−.412	.167	−.207	.152	−.107	.155	−.342	.203
Diet index	.087	.072	.075	.076	−.011	.059	.063	.060	.125	.078
Alcohol	.019	.082	.010	.067	−.308	.148	−.134	.148	−.198	.191
Acculturation	.095	.120	.209	.127	.111	.117	.261	.119	.183	.154
Education	.053	.023	.018	.025	.033	.022	.026	.022	.010	.029
Income	.005	.011	.031	.012	−.014	.011	.012	.011	.031	.015
Age	.020	.007	.002	.007	.007	.006	.000	.005	−.014	.008
Puerto Ricans										
Smoking	−.087	.146	−.084	.168	−.056	.132	.043	.148	.218	.178
Diet index	.053	.071	.057	.052	.057	.064	−.012	.072	.138	.087
Alcohol	−.020	.076	−.145	.079	−.073	.096	−.008	.108	−.122	.131
Acculturation	.182	.130	.074	.149	−.086	.149	−.167	.164	−.171	.199
Education	.053	.025	.048	.029	.034	.025	.063	.029	.031	.035
Income	−.006	.013	.016	.015	.001	.016	.010	.018	.015	.022
Age	.022	.009	.006	.010	.007	.004	−.004	.005	.008	.006

Entries are weighted regression coefficients and adjusted standard errors. The health service measures were coded with higher numbers indicating more recent use of the service. Cigarette smoking was a dichotomous variable with current smoker = 1 and nonsmoker = 0.

TABLE 5—Weighted Prevalence of Recent Screening by Health Risk Behaviors

	Males		Females		
Recency:	Physical 0–2 yrs	Dental 0–2 yrs	Physical 0–2 yrs	Dental 0–2 yrs	Pap 0–2 yrs
Mexican Americans					
Cigarette smoking					
Smoker	29.7	42.1	42.3	57.3	75.6
Nonsmoker	37.4	55.9	47.3	62.2	76.0
Diet[a]					
Low	31.0	46.3	42.7	55.0	72.8
High	37.0	53.6	49.3	67.0	79.2
Alcohol use					
Moderate/high	30.8	52.2	45.2*	70.0*	83.0*
None/low	36.7	47.9	46.2	60.3	75.3
Cuban Americans					
Cigarette smoking					
Smoker	34.6	51.8	54.1	70.6	59.4
Nonsmoker	49.8	71.4	60.2	73.6	71.5
Diet					
Low	38.9	59.3	57.3	71.7	66.9
High	46.9	65.8	60.2	74.0	70.4
Alcohol use					
Moderate/high	41.1	60.2	61.4*	70.0*	89.4*
None/low	44.1	64.0	58.7	73.0	67.7
Puerto Ricans					
Cigarette smoking					
Smoker	45.7	59.8	59.0	74.9	80.5
Nonsmoker	49.3	63.5	63.5	74.9	78.8
Diet					
Low	46.6	56.9	55.9	72.9	76.0
High	48.6	66.0	67.6	76.7	82.6
Alcohol use					
Moderate/high	44.9	58.4	54.3*	81.1*	76.5*
None/low	49.9	64.5	62.8	74.3	79.6

Entries are not age-adjusted.

*An asterisk next to an entry indicates that the statistic should be viewed cautiously because it does not meet the minimum reliability standard due to small cell size.

[a]The "low" diet score includes individuals who received a score of 0 or 1 on the overall diet index; the "high" diet score includes persons who received a score of 2 or 3 on the overall diet index.

association between dietary behavior and screening was stronger for the women than the men, the relation between smoking and screening was stronger for the men than the women. These gender-specific associations were not as clear for the Cuban Americans or Puerto Ricans.

Alcohol—The results for alcohol consumption were mixed. Some of the regression coefficients suggested that heavier drinking tended to associate with less recent use of the health services (e.g., Mexican American men/physical examination), while other coefficients suggested the opposite. We attempted to clarify the picture by examining a subset of heavy drinkers who did not suffer from chronic illnesses (diabetes, heart disease, hypertension) that may necessitate regular visits to a physician. Analyses of this subset, however, yielded essentially the same null and mixed results. Second, among those with no chronic illnesses we compared screening rates for two extreme groups: nondrinkers versus heavy drinkers (those drinking an average of 1.00 or more oz of ethanol per day). Again, most of the results were nonsignificant. Thus, the hypothesis regarding the association between heavy drinking and infrequent use of preventive health services was not confirmed.

Acculturation and Sociodemographics—These factors did not produce strong effects. Increased acculturation tended to be associated with greater recency of screening for the Mexican Americans and Cuban Americans. Higher levels of education and higher income (primarily for the Mexican American men and women) were related to more recent use of the services, although the regression weights were small. The effects of age were minimal.

Discussion

Cigarette smoking, poor dietary practices, and excessive alcohol consumption place individuals at risk of specific disease.[20–27] Ideally, those who engage in these health risk behaviors should use preventive health services regularly in order to monitor their health status. The present study demonstrated several instances in which those at risk were screened *less* recently than those not at risk. Such underutilization increases the likelihood of diagnosis of illness at a later stage.

For each sample, risk of disease was greater for men than women. Men were more likely to be smokers, heavier smokers, and heavier alcohol users; also, their screening rates were much lower than the women's. These gender differences held across level of acculturation, which suggests that similar results would be found in non-Hispanic populations.

The gender effects may stem from sex-role norms. The strong emphasis in our society on self-reliance and independence for men may contribute to their less-than-optimal health attitudes and practices.[33] Further, these norms may influence personal beliefs about one's invulnerability to illness,[34] which may be held more strongly by Hispanic men than Hispanic women. Such beliefs are likely to impact negatively on preventive action.[35]

We found significant positive correlations between acculturation and alcohol consumption, especially for the women of each sample. These findings are consistent with the acculturation model of alcohol use,[36–38] which emphasizes normative influence on behavior. The dominant norms guiding alcohol consumption among non-Hispanic women are more liberal than the cultural norms for Hispanic women—non-Hispanic women drink more frequently and drink greater quantities of alcohol than Hispanic women in general.[36] As Hispanic women acculturate, their drinking increases because their behavior becomes more strongly influenced by the norms and practices of the dominant group. The smaller correlation between acculturation and alcohol use for Hispanic men may reflect the fact that, although the specific pattern of frequency versus quantity of alcohol consumed differs somewhat between Hispanic and non-Hispanic males,[36] the drinking norms for these two groups are fairly similar.

Health risk behaviors such as cigarette smoking and poor dietary practices were associated with less recent screening activity. In particular, for the Mexican American and Cuban American men, cigarette smoking correlated strongly with less recent utilization of routine dental and physical examinations; for the women especially the Mexican American women), poor dietary practices correlated strongly with less recent screening, including Pap smear. Because the HHANES was a cross-sectional design, we cannot conclude that cigarette smoking or poor dietary practices causes one to use screening services infrequently. Rather, the data suggest that these behaviors tend to co-occur. At the outset, we suggested that the relationships may stem from a common factor, namely one's underlying attitude toward health promotion and disease prevention. It is also quite possible, however, that some smokers may actively avoid check-ups from doctors because of their anxiety about receiving negative feedback about their health or being told by the physician to quit smoking. Alternatively, some smokers may still be unaware of the link between smoking and disease, or may discount or deny that link, and thus use screening services infrequently.[39] These possibilities are not limited to any one subpopulation. The fact that we found the same inverse association between smoking and utilization for high and low acculturated Mexican American and Cuban American men suggests that our findings would generalize to non-Hispanic White populations. We should note, however, that although the associations may hold for each of these groups, the absolute level of medical screening is lower among Hispanics.[1,3]

Finally, alcohol consumption was not strongly or consistently related to screening in any of the three samples. Unfortunately, the present data do not permit a conclusion regarding the role of this health risk behavior in screening activity. One potentially important factor, however, is a person's perception of the health effects of alcohol use, which is likely to vary considerably among individuals. For example, some people with a relatively positive attitude toward health promotion may drink moderately/heavily due to the fact that they perceive the risk of drinking to be slight. Others may perceive the risk of drinking differently. This variability may occur because, compared to the clear media message about the negative effects of smoking cigarettes, the health consequences of alcohol have not been communicated as clearly or consistently. The variable beliefs about the effects of alcohol use may have contributed to the inconsistent findings for alcohol consumption and medical screening.

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IV. Health Care Utilization Barriers among Mexican Americans: Evidence from HHANES 1982–84

Antonio L. Estrada, PhD, MSPH, Fernando M. Treviño, PhD, MPH, and Laura A. Ray, MPA

Abstract: Data from the Hispanic Health and Nutrition Examination Survey (HHANES) conducted by the National Center for Health Statistics in 1982–1984 were analyzed to document the type of barriers encountered which prevented Mexican Americans from obtaining health care, the sociodemographic subgroups most vulnerable to such barriers, and to examine the combined effects of predisposing, enabling, and need characteristics on these barriers.

The findings suggest, in general, that low income groups, younger age groups, the less acculturated, those who lack health insurance coverage, those with functional limitations, and those in poorer perceived health status encounter more barriers than others, and are prevented by these barriers from obtaining health care for themselves. [*Am J Public Health* 1990; 80(Suppl):27–31.]

Introduction

Mexican Americans have been found to make less use of health care than White and Black non-Hispanics, Puerto Ricans, Cuban Americans and other Hispanics.[1,2] Previous research has identified various potential barriers which may impede Mexican Americans from obtaining medical care. These barriers include, but are not limited to, language and cultural differences, lack of transportation, geographic inaccessibility, financial constraints such as the cost of health care and limited health insurance coverage, and isolation from the mainstream culture.[3–9]

Several models have been proposed to examine the utilization of health services.[10–14] One of the most widely used models is a behavioral model developed by Andersen[14] and subsequently expanded by Andersen and Newman,[15] Aday and Andersen,[16] and Andersen, *et al.*[8] In its original formulation Andersen's model conceived of health care utilization as a consequence of three general factors: 1) predisposing characteristics such as age and gender; 2) enabling characteristics such as income and health insurance coverage; and 3) need characteristics such as health status. The components of this model were hypothesized to contribute independently to the understanding of differential utilization of health services. Need characteristics have been found to be more important in predicting utilization of health services than either predisposing or enabling characteristics.[17–19]

Mechanic,[17] Rundall,[18] and Wolinsky[19] have noted several problems with the model in its current form: the low amount of utilization variance explained by the model; the specification of the model in practice as additive, when in fact the model as stated by Andersen and colleagues suggests characteristics, over and above that explained by predisposing and enabling characteristics.

The approach taken in this paper uses a multivariate model based upon Andersen and colleagues' behavioral model of health care utilization.[7,8,14,16] This model focuses on utilization barriers using health care utilization as a control in order to examine variables related to encountering barriers and being prevented by these barriers from obtaining health care.

The over-riding rationale in the model's application to utilization barriers is that if predisposing, enabling, and need characteristics help explain the over- or underutilization of health care, these same constructs may help explain why some people encounter barriers and are thus prevented from utilizing health care services.

The purpose of the present study is three-fold: to document the type of barriers to health care reported by Mexican Americans; to determine which sociodemographic subgroups are most vulnerable to such barriers; and to examine the joint effects of predisposing, enabling, and need characteristics on the barriers to health care.

Methods

In the analyses presented in this paper, sampling weights were employed to arrive at reliable point estimates, and corrections for the clustering effect introduced by the complex sample design were accounted for when means and standard errors were computed by a method developed at the National Center for Health Statistics[20] and documented in this issue of the Journal.[21] The current analyses focus on 3,935 Mexican Americans ages 20–74 years who were eligible for inclusion based on their having self-identified themselves as either Mexican, Mexican American, or Chicano.

In keeping with Andersen's multivariate framework, the present study includes predisposing, enabling, and need characteristics. Table 1 displays each of the variables under these constructs differentiated by gender and age group as well as the percentage of persons who reported encountering barriers and the percentage of persons who reported having had a physician visit within the past year.

Dependent Variables

Persons interviewed in the Hispanic Health and Nutrition Examination Survey (HHANES) were asked if they encountered a barrier when they last attempted to utilize health care (from a list of 13 possible barriers, Table 2). If they reported encountering a barrier, they were further asked if this barrier actually prevented them from receiving health care.

The two major dependent variables of interest therefore were the extent to which respondents encountered utilization barriers and whether or not these barriers prevented them from obtaining care for themselves. Utilization of a physician within the past year was used as a control variable when each of the dependent variables was examined. Barriers which prevented care from being received were dichotomized as "prevented" and "not prevented" from obtaining health care. Utilization barriers encountered is a continuous variable derived by dividing the sum of those barriers the respondent reported encountering by the total possible number of barriers reported. Not all persons could report en-

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NOTE: Author affiliations are listed elsewhere under CONTRIBUTORS.

TABLE 1—Sociodemographic and Health Utilization Related Variables by Gender and Age Group

	Ages 20–30		Ages 31–45		Ages 46–74	
Characteristics	Male (N = 628)	Female (N = 696)	Male (N = 553)	Female (N = 646)	Male (N = 616)	Female (N = 796)
Predisposing						
Mean age	24.8	24.9	36.6	36.9	56.9	57.1
Mean Aculturation	2.6	2.7	2.6	2.6	2.4	2.3
% Married	60.1	65.2	85.9	73.9	84.5	63.8
% US born	59.3	65.7	59.6	62.3	64.4	67.3
Enabling						
Mean income	15.8	15.1	17.5	15.9	15.1	13.4
Mean education	10.6	10.3	9.9	9.5	7.3	6.7
% In poverty	23.7	30.1	19.6	29.6	27.4	36.5
% Employed	92.3	52.4	95.6	54.9	65.8	34.8
% Speak english	61.5	65.3	63.4	64.6	55.6	52.9
% Have regular source of care	55.9	77.6	68.6	82.9	78.1	86.0
Type of Insurance:						
% Not covered	43.8	38.1	29.6	31.6	26.5	33.2
% Private	54.1	53.4	66.9	58.8	47.5	41.3
% Medicare	.3	.8	1.0	1.3	18.0	13.2
% Medicaid	1.3	6.9	.8	6.6	6.2	10.4
% Other public	.4	.9	1.7	1.8	1.8	1.9
Need						
% Impaired	43.3	4.4	4.6	8.2	17.6	18.9
Perceived health:						
% Poor	2.1	2.2	2.9	4.0	11.5	15.0
% Fair	22.4	24.5	23.3	32.8	35.4	39.8
% Good	37.0	38.9	36.4	32.9	31.6	28.8
% Very good	20.2	18.9	21.1	17.7	13.9	10.4
% Excellent	18.2	15.5	16.3	12.6	7.6	6.0
Outcome						
Mean % barriers	7.0	8.8	5.4	7.4	4.2	7.5
% Last physician visit <1 year ago	51.0	80.0	56.8	72.2	63.2	76.7

TABLE 2—Rank Order of Specific Barrier Items, Mexican Americans (N = 3935)

Barriers Encountered Which Impeded Care	Persons Reported Encountering Barrier(s) N	Persons Reported Encountering Barrier(s) %	Persons Encountered Barrier(s) Reported Care Prevented* %
Cost of health care	706	18.0	82.3
Had to wait too long in the office or clinic	376	9.5	58.9
Had to wait too long to get an appointment	365	9.3	67.0
Would lose pay from work (applicable only to those who work)	337	8.6	72.7
The hours were inconvenient	307	7.8	70.3
Care was not available when needed	256	6.5	74.0
Did not know where to go	228	5.8	75.0
Did not have transportation	224	5.7	74.2
Had no confidence in the staff	168	4.3	62.9
Needed someone to take care of your children (applicable only to those with children)	132	4.2	78.3
Staff did not speak Spanish	155	3.9	62.8
Staff was disrespectful	109	2.8	53.2
There were no Hispanic staff members	91	2.3	65.6

*The percent reported for prevented barriers is derived from those persons who acknowledged an encountered barrier. For example, 82.3 percent of those who encountered cost as a barrier said it actually prevented them from obtaining care for themselves.

countering all 13 barriers. Persons who were employed and had children could report all 13 possible barriers. Persons who either were not employed or did not have children could only report 12 possible barriers. Persons who were unemployed and had no children could only legitimately report 11 of the possible 13 barriers on the list.

Results

Sixty-seven percent of the Mexican American sample did not report encountering any barriers while approximately 33 percent of the sample reported encountering one or more barriers when last attempting to obtain health care. Among those who reported encountering barriers to health care, 73 percent reported that the barrier(s) were serious enough to actually prevent them from obtaining health care on that occasion.

The encountered and prevented barrier items were rank-ordered by the percent of respondents who reported each item (Table 2). The five most commonly encountered barriers reflect dimensions of availability and accessibility of health care. For example, cost of health care was the most frequently encountered barrier among Mexican Americans, followed by having to wait too long at the doctor's office or clinic, waiting too long to get an appointment to be seen, inconvenient clinic or office hours, and losing income as a result of having to be away from work to obtain health care.

The five barriers which were most frequently reported as actually preventing Mexican Americans from obtaining health care were somewhat different. While cost of health care was the most frequently reported, needing someone to take care of their children while they obtained health care was the second most cited barrier which prevented them from obtaining health care, followed by not having transportation to get to the source of health care, and losing pay from work while obtaining care.

In general, relatively few Mexican Americans reported encountering barriers which reflect the cultural and linguistic aspects of health care (e.g. no Spanish-speaking staff members, no Hispanic staff members, staff being disrespectful). The data seem to suggest that cost, availability, and accessibility constraints were more important than cultural constraints as measured by the HHANES.

Bivariate Associations

Predisposing Characteristics—The relationships between selected predisposing characteristics and utilization barriers were examined so as to identify sociodemographic subgroups within the Mexican American population who were most vulnerable to encountering barriers to health care. Results revealed that Mexican Americans who are younger in age, female, less acculturated, and foreign born were more likely to be prevented from obtaining health care than others.

Enabling Characteristics—Overall, the enabling characteristics examined showed a consistent and significant association with utilization barriers. Persons with low incomes or low educational levels, Spanish-speakers, those with incomes below the poverty level, the unemployed, the uninsured, and those without a regular source of care reported utilizing health care services less frequently and were more likely to encounter barriers to obtaining health care. These variables were interrelated: e.g. the unemployed have lower incomes and probably have less insurance coverage than the employed. The predisposing and enabling characteristics suggest that a significant proportion of those who can least afford health care and who are least equipped to deal with the American health care system were prevented from using health care services.

Need Characteristics—Both need characteristics examined (perceived health status and functional impairment) were found to be positively associated with encountering barriers and thus being prevented from obtaining health care. Obviously, persons in poor health and those with functional impairments have greater opportunity to encounter barriers to care due to their greater need for care and greater utilization of health services. Nonetheless, Mexican Americans with greater health needs reported encountering more barriers.

Multivariate Analyses

To assess the combined effects of predisposing, enabling, and need characteristics, separate multivariate analyses were performed with utilization barriers encountered and prevented as the dependent variables, respectively, controlling for time since last physician visit. All independent variables were entered simultaneously into each multivariate model. Therefore, the coefficients reported in Tables 3 an 4 estimate the relationship of each independent variable on the dependent variable, while simultaneously taking into account all other independent variables.

The control variable, time since last physician utilization, was dichotomized as "0" meaning less than one year and "1" meaning greater than one year. Barriers which were reported to have prevented a person from obtaining care were dichotomized as "0" meaning no barriers and "1" meaning one or more barriers were reported. Design effects were

TABLE 3—Results of Multiple Regression for Dependent Variable: Percent Barriers Encountered by Age and Gender

	Visited Physician within Last Year				Visited Physician One Year or More Ago			
	Males		Females		Males		Females	
Variables	b	SE	b	SE	b	SE	b	SE
Ages 20 to 30								
Acculturation	.0098	.013	.0033	.008	−.0034	.013	−.0146	.017
Marital status	−.0067	.003	.0023	.003	−.0048	.003	.0037	.007
Income	−.0019	.001	−.0032	.001	.0001	.001	−.0062	.002
Education	.0015	.003	.0004	.003	−.0000	.003	.0024	.007
Have regular source of care	−.0265	.018	−.0439	.018	.0050	.018	−.0258	.004
Type of insurance	−.0166	.019	.0045	.008	−.0[illegible]58	.021	.0215	.025
Impaired	.0200	.053	.0210	.031	.0307	.059	.0007	.085
Perceived health	−.0365	.010	−.0169	.007	−.0165	.010	.0024	.008
Ages 31 to 45								
Acculturation	−.0065	.011	−.0129	.010	.0144	.010	.0105	.014
Marital status	.0117	.006	.0040	.005	.0088	.006	−.0046	.006
Income	−.0014	.001	−.0010	.001	−.0034	.002	−.0034	.002
Education	−.0018	.003	−.0010	.003	.0039	.002	−.0017	.004
Have regular source of care	−.0004	.022	.0240	.024	−.0304	.017	−.0196	.023
Type of insurance	−.0191	.013	−.0055	.009	.0019	.014	.0032	.015
Impaired	.0063	.041	.0068	.025	.0184	.047	−.0186	.069
Perceived health	−.0031	.010	−.0178	.007	−.0196	.010	−.0214	.013
Ages 46 and Older								
Acculturation	−.0044	.007	−.0196	.008	−.0139	.008	.0077	.015
Marital status	−.0062	.005	−.0009	.004	−.0012	.006	.0291	.007
Income	.0009	.009	−.0019	.001	.0003	.001	.0023	.002
Education	−.0036	.002	−.0006	.002	.0005	.002	−.0086	.003
Have regular source of care	−.0407	.019	.0154	.024	.0207	.014	−.0079	.024
Type of insurance	−.0106	.007	−.0222	.006	−.0137	.008	−.0117	.013
Impaired	.0302	.015	.0922	.016	.0215	.027	.0366	.034
Perceived health	−.0052	.006	−.0106	.007	−.0078	.006	−.0206	.014

TABLE 4—Multiple Logistic Regression Coefficients Estimating the Effect of Predisposing, Enabling, and Need Characteristics on Barriers Prevented Care Those Reporting They Visited the Physician within the Last Year (0 to 11 Months) (N = 793)

Variables	Regression Coefficient	S.E.	Odds Ratio	(Confidence Intervals)
Those Reporting Visit to Physician within the Last Year (N = 793)				
Age	−.0076	.008	.98	(.95, 1.02)
Gender (1 = Male, 2 = Female)	.1220	.167	1.32	(.62, 2.81)
Marital status (1 = married, 2 = not married)	−.0595	.043	.87	(.72, 1.06)
Annual income	−.0457	.013	.90	(.85, .95)
Education level	.0083	.027	1.02	(.90, 1.15)
Acculturation level	−.0792	.098	.92	(.45, 1.30)
Perceived health status	−.1462	.082	.71	(.49, 1.03)
Impairments (1 = yes, 2 = no)	.5828	.309	3.82	(.84, 15.43)
Type of insurance	−.1555	.096	.70	(.45, 1.08)
Regular source of care (1 = yes, 2 = no)	−.3400	.235	.46	(.16, 1.32)
Those Reporting Visit to Physician Longer than One Year Ago (N = 304)				
Age	−.0413	.014	.91	(.85, .97)
Gender (1 = Male, 2 = Female)	.5262	.311	3.36	(.83, 13.66)
Marital status (1 = married, 2 = not married)	.1298	.084	1.35	(.92, 1.97)
Annual income	−.0178	.024	.96	(.86, 1.07)
Education level	−.0498	.048	.89	(.88, .90)
Acculturation level	.1014	.173	1.26	(.58, 2.76)
Perceived health status	−.3404	.150	.46	(.23, .90)
Impairments (1 = Yes, 2 = No)	.3500	.663	2.24	(.11, 44.61)
Type of insurance	−.0999	.207	.80	(.31, 2.02)
Regular source of care (1 = yes, 2 = no)	−.1249	.320	1.33	(.31, 5.65)

Original sample size is 1109 because only those who reported encountering a barrier are included. Controlling for visit greater than one year and listwise deletion of missing variables reduces the N to 304.

Standard errors adjusted when design effect >1.

computed using SESUDAAN, which incorporates the Taylor Approximation Method. Standard errors were then adjusted when design effects were greater than one.

Utilization Barriers Encountered as Dependent Variable—Regressions were run separately for each age-gender group using utilization of health services in the past year as a control variable (see Table 3).

Among persons who have visited a physician within the past year, the strongest variable associated with encountering barriers among Mexican American men ages 20–45 years was marital status, with married men encountering more barriers to health care. Among Mexican American men ages 46–74 years, those with less years of formal education, those with no regular source of care, and those with functional impairments encountered more barriers than others. For Mexican American women ages 20–45 years, poorer perceived health status, low income, and not having a regular source of care were associated with encountering barriers when seeking care. Among Mexican American women 46–74 years of age, the less acculturated, the uninsured, and those with functional impairments encountered more barriers than others.

For Mexican American men ages 20–30 years who have not visited a physician for one year or more, lack of health insurance coverage and poorer perceived health status were associated with encountering barriers. There were no associations of consequence for Mexican American men 31 years of age and older. For Mexican American women ages 20–30 years, those with no health insurance coverage and those in poorer perceived health status encountered more barriers than others. Among Mexican American women ages 31–45 years, those with low incomes and those in poorer perceived health status encountered more barriers; for Mexican American women ages 46–74 years, being married and having less years of education were associated with encountering barriers when seeking health care.

Barriers Which Prevented Health Care as Dependent Variable—Multiple logistic regression analysis was the method of choice employed to examine the contributions of the predisposing, enabling, and need variables because of the skewed distribution and dichotomy of the dependent variable. Utilization of health services in the past year was again used as a control variable. Table 4 displays the results of this analysis which included only persons who reported encountering a utilization barrier.

Results indicated that, among Mexican Americans who saw a physician within the past year, those with lower incomes and those with functional impairments were more likely to be prevented from obtaining care than others. Among Mexican Americans who had not visited a physician within the past year, younger persons, women, and those in poorer perceived health status were more likely than others to have been prevented from obtaining health care.

Discussion

Findings from the HHANES revealed that approximately one-third of the adult Mexican American population had reported barriers to obtaining health care during their most recent medical encounter. These barriers to receiving health care were significant in that three-fourths of the time the barriers encountered were sufficient to prevent persons from obtaining the medical care they sought.

Analyses revealed that Mexican Americans with low educational levels and low incomes are experiencing more problems than other Mexican Americans in accessing health care. Spanish-speaking, foreign-born, and less acculturated Mexican Americans were found to be the most likely to be prevented from obtaining the care they sought.

A significant minority of the adult Mexican American population reported encountering barriers which prevented them from accessing medical care. Almost one in five adult

Mexican Americans reported the cost of medical care was a barrier to their receiving care. The concern over the cost of health care among Mexican Americans would appear to be warranted, given that they have the highest rate of being uninsured for medical expenses in the country.[2] A smaller proportion of the adult Mexican American population reported problems accessing medical care due to the lack of health care services during nonworking hours (e.g. losing pay from work, hours of operation were inconvenient, care was not available when needed). Occupational analyses were not performed; however, it may be that access to medical care is more dependent upon the availability of expanded hours of operation among health providers due to the greater propensity of Mexican Americans to be employed in hourly-paid service occupations.

Given the shortage of minority and multilingual health providers in the United States, it was encouraging to find that fewer Mexican Americans reported cultural or linguistic barriers to care (e.g., no Spanish-speaking staff, no Hispanic staff, disrespectful staff). This finding was unexpected given the body of research which has suggested that the exclusion of Spanish-speaking and/or Hispanic staff members is a primary reason for underutilization of health services among Mexican Americans.[4,6,22–24] This finding does not refute the probable importance of cultural and linguistic aspects of medical care but rather may signify that among Mexican Americans the cost, availability and accessibility constraints may be more important in determining utilization of health services than the cultural and linguistic constraints (as measured in the HHANES). It is possible that accessibility and availability of care (cost, geographic location, hours of operation, etc.) are most closely related to utilization of services while cultural and linguistic aspects of care relate more to satisfaction with care, compliance with treatment, continuation in treatment, etc.

It is unfortunate that Mexican Americans with the highest need for health care (poor perceived health status and functional limitations) tended to report the highest rates of encountering barriers which prevented them from obtaining the needed care. Also, approximately 6 percent of the adult Mexican American population reported not knowing where to go for health services or having a way to get there. Health providers and agencies seeking to increase access to medical care among Mexican Americans can assist by focusing on service delivery aspects (e.g., cost of care, providing convenient hours for the community they serve, decreased waiting time, etc.) and outreach aspects (educational and transportation) to reach those individuals who may not know where to go for care. It is possible that such efforts would increase access to care for all ethnic/racial groups.

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V. Utilization of Curanderos by Mexican Americans: Prevalence and Predictors Findings from HHANES 1982–84

JOHN C. HIGGINBOTHAM, MPH, FERNANDO M. TREVIÑO, PHD, MPH, AND LAURA A. RAY, MPA

Abstract: Data from the Southwest sample of the Hispanic Health and Nutrition Examination Survey (HHANES) were analyzed to examine whether the use of a curandero or other folk medicine practitioner hindered, enhanced, or did not affect the utilization of western health care services by Mexican Americans. Findings revealed that only 4.2 percent of the HHANES sample persons between the ages of 18–74 years reported consulting a curandero, herbalista, or other folk medicine practitioner within the 12 months prior to the survey. Income, self-perceived health status, the language of the interview, and dissatisfaction with modern medical care recently received independently predicted curandero utilization (adjusted OR 2.01 and 1.66, respectively). Low income and self-perceived health status were less strongly related to curandero utilization. [*Am J Public Health* 1990; 80(Suppl):32–35.]

Introduction

Mexican Americans have long been thought to be in generally poorer health than White non-Hispanics as evidenced by shorter life expectancies, and higher infectious, parasitic disease, infant and maternal mortality rates; they appear to be in better health than Black non-Hispanics whose socioeconomic conditions are generally comparable.[1,2] Recent evidence indicates that these disadvantages are diminishing.[3–6] Less progress has been made, however, for those conditions usually associated with a lack of access to health care, such as tuberculosis and diabetes.[2,7–9]

Despite their apparent health needs, Mexican Americans have not achieved expected levels of health care utilization. Unlike Puerto Ricans and Cuban Americans who exhibit higher levels of health care utilization than other White Americans, Mexican Americans demonstrate the lowest level of health care utilization of all ethnic/racial groups (for whom there are data).[6] This underutilization of modern medical care has occurred despite factors which in other groups have been indicative of higher utilization rates (e.g., the stress factors of poverty, acculturation, and crowded housing).[10]

Alternative hypotheses to account for this underutilization include:

- cultural and language differences between patient and provider;
- folk medicine practices as an alternative to that of the modern health care system;
- environmental and socioeconomic conditions creating priorities for daily living inconsistent with the fee-for-service modern health care system.[11]

Contradictory results have been reported on why Mexican Americans underutilize health care services. Some of this research, particularly concerning the use of folk medicine, has been considered a disservice to Hispanics because it focuses on cultural or attitudinal explanations for underutilization while ignoring other possible explanations such as those associated with socioeconomic factors and lack of access to care.[12]

Andersen, *et al*, posited that Hispanics have limited access to medical care due to low levels of income, education, and health insurance coverage.[13] Treviño and Moss found that Mexican Americans with health insurance are 1.5 times more likely to consult a physician than Mexican Americans without insurance.[14] In addition to the monetary restraints, Chesney, *et al*, suggested that acculturation and social isolation are pertinent barriers to medical care for Mexican Americans.[15] While the debate continues as to what the exact barriers to health care are for Mexican Americans, the general consensus is that for Mexican Americans the availability of health care (as distinct from access) is comparable to that of the population in general.[13]

All societies have some systematized manner in which to deal with illness.[16] In the past some Mexican Americans have utilized curanderos or folk medicine practitioners. The number of Mexican Americans who still use curanderos as an alternative to modern medical care, for cultural and/or economic reasons, is unknown.

Most studies of curanderismo, the practice of folk medicine, have sought to describe and evaluate this phenomenon by indirect methods, i.e., through studies focused on the practitioner's patients.[17,18] These studies have produced detailed descriptions of the specific healing techniques used[19] or information on the type of disorder being treated.[17] Little is known about the prevalence of their use by Mexican Americans and what factors differentiate users from nonusers. Such factors could include cultural orientation, ability to pay, and dissatisfaction with previous medical care. It is unknown whether curanderismo is used as an exclusive alternative to modern medicine or whether both types of care are used concurrently.

This study seeks to compare those Mexican Americans (ages 18 years and older) in the Southwestern sample of the HHANES who reported consulting a curandero, herbalista, or other folk medicine practitioner within the last 12 months with those who did not do so. A model to predict those who sought alternative care is formulated.

Methods

The Hispanic Health and Nutrition Examination Survey (HHANES), the first national study of the health of the Hispanic population, was conducted during 1982–84 by the National Center for Health Statistics. A more detailed account of the design is presented elsewhere.[20,21] The HHANES sought to obtain representative samples of Hispanics households by concentrating sampling procedures for Mexican Americans in five southwestern States.

Specifically, this analysis includes those Mexican Americans, ages 18 years and older, who responded to the question

NOTE: Author affiliations are listed elsewhere under CONTRIBUTORS.

"There are some providers of health care that we sometimes go to, such as curanderos, sobadores, herbalists, spiritualists, and others. Have you seen or talked to any of these persons for health care during the past 12 months?" There are significant differences between curanderos, herbalistas, and other folk medicine practitioners; however, all were included in the single question used in HHANES to access folk medicine utilization. Further, they all represent a traditional medicine alternative to "scientific" medicine. Henceforth in this report, curandero will be used to refer to all of these folk medicine practitioners.

Due to the nature of these data, efforts were undertaken to account for the large size and method of sampling in the analysis. Design effects have been accounted for through the reporting of weighted percentages and adjusted variances, which were computed utilizing the Taylor approximation of the SESUDAAN methodology.[22] Comparisons were made between weighted and unweighted estimates resulting in insignificant variation, sample sizes are presented unweighted while statistical test and percentages utilize weighted estimates.

The variables used in the prediction model were drawn from the literature. They included indicators of: health status, ability to pay for health care, and health care utilization (see Appendix for details of these items), as well as the demographic variables of gender, age, marital status, education, nativity, acculturation, language of interview, family size, and size of the respondent's community or city of residence.

Descriptive and multivariate analyses, including logistic regression (using SAS, Proc Logist; SAS Institute, Inc.), were performed on the data in order to build the model which best predicted curandero utilization with the fewest variables. Odds ratios and 95 percent confidence intervals, using adjusted standard errors, were calculated from beta coefficients.[23]

Results

The total sample size for this investigation includes 3,623 Mexican Americans from the Southwestern United States ages 18 years and over. Of this number, only 148 persons (4.2 percent) reported consulting a curandero, herbalista, or other folk medicine practitioner within the last 12 months.

Users of curanderos were slightly more likely to be male, less well educated, and foreign born (Table 1). Marital status, age, family size, geographical distribution, or size of the community in which they reside were not found to be related to the use of curanderos.

Acculturation was not found to be significantly related to use of curanderos. Subsequent analysis, however, revealed that being interviewed in Spanish was found to be highly predictive of curandero utilization, whereas nativity was not.

Table 2 presents the comparisons of proportions on indicators of health status by users and non-users of curanderismo. During the physical examination portion of the HHANES, the physician was asked to indicate his/her global assessment of the respondent's health status. The vast majority (>85 percent) for both groups were classified by the physician as being in very good or excellent health. There were no differences in physician assessment of health status of Mexican Americans who used curanderos and those who did not utilize them.

Those who consulted a curandero were more likely than those who did not to report their current perceived health status as fair or poor and to believe they had little control over their future health status.

TABLE 1—Demographic Characteristics of Reported Users and Non-Users of Curanderismo

	% Users	% Non-Users
(N)	(148)	(3475)
Gender		
Male	58.1	50.0
Female	41.9	50.0
Ages (years)		
18–44	64.2	65.2
45–64	31.1	29.2
65–74	4.7	5.6
Marital Status		
Married: spouse in household	67.1	66.3
Married: spouse not in household	3.3	1.9
Widowed	5.4	3.5
Divorced	5.3	5.5
Separated	3.3	3.9
Never Married	15.6	19.0
Education		
0–8 years	46.6	36.6
9–11 years	15.4	20.0
12+ years	37.9	43.5
Nativity		
United States	52.0	61.0
Other	48.0	39.0
Acculturation		
1.0–1.9	48.4	33.9
2.0–2.9	17.6	21.1
3.0–3.9	31.1	38.5
4.0–4.9	2.9	6.4
Language of Interview		
English	48.6	61.3
Spanish	51.4	38.7
Family Size (persons)		
1 or 2	22.2	19.6
3 or 4	39.6	40.5
5 or 6	25.3	27.6
7 or more	12.9	12.4
Size of Place (population)		
500,000 or more	27.9	24.2
100,000–499,999	14.4	15.0
25,000–99,999	28.2	27.5
200–24,999	21.7	24.1
Not in a place	7.9	9.2

Those who consulted curanderos in the last year were found to be less satisfied with the care they received at their last visit to a medical practitioner. However, no differences were found regarding having a regular source of care or recency of last visit to a medical practitioner (Table 3). Further, poverty status, the presence or absence of health insurance coverage or the type of coverage were not found to be related to use of curanderos (Table 4).

Logistic regression with backward elimination was applied to this sample in order to delineate which variables out of this theoretical framework were most predictive of curandero utilization. Those variables that remained within the model were: satisfaction with care received at last visit, the language of the interview, self-perceived health status and income (Table 5).

Discussion

As noted by others,[16] relatively few persons of Mexican origin in the United States consult a curandero. Only 4 percent of the adult Mexican Americans interviewed in the HHANES reported that they had consulted a curandero within the past 12 months. Comparable population estimates

TABLE 2—Indicators* of Health Status by Reported Users and Non-Users of Curanderismo

	% Users	% Non-Users
(N)	(148)	(3475)
Health Status (Physician Assessment)		
Excellent	52.9	59.4
Very Good	36.5	28.0
Good	9.0	9.2
Fair	1.1	2.8
Poor	0.5	0.5
Health Status (Self-Assessment)		
Excellent	11.8	13.8
Very Good	15.5	18.1
Good	30.7	34.7
Fair	31.7	28.3
Poor	10.4	5.1
Future Health Locus of Control		
Excellent	37.7	43.1
Good	43.4	44.6
Fair	17.1	8.9
Poor	1.8	3.4

*See Appendix for definitions.

TABLE 3—Indicators* of Health Care Utilization by Reported Users and Non-Users of Curanderismo

	% Users	% Non-Users
(N)	(148)	(3475)
Source of Care		
Regular only	58.8	66.2
Regular & other	7.4	7.1
Other	29.7	24.6
Never visited	4.1	2.1
Visit Interval		
Never	4.1	2.1
<1 month	19.9	19.3
>1 & <12	50.2	47.7
>1 year & <5	23.4	25.0
>5 years	3.3	6.0
Satisfaction with Care		
Very satisfied	69.2	80.8
Somewhat satisfied	14.2	14.4
Not at all satisfied	16.6	4.7

*See Appendix for definitions.

TABLE 4—Indicators* of Ability to Pay for Health Care Services by Reported Users and Non-Users of Curanderismo

	% Users	% Non-Users
Income		
<$20,000	36.7	35.0
≥ $20,000	63.3	65.0
Type of Health Insurance		
Not covered	39.0	35.3
Private	50.2	53.9
Medicare	4.3	4.4
Medicaid	4.9	5.0
Other public	1.7	1.3
Poverty Level		
At or below	24.2	28.1
Above	75.8	71.9

*See Appendix for definitions.

TABLE 5—Logistic Regression Odds Ratios (OR) and 95% Confidence Interval (CI) for the Association of Curandero Utilization with Predictor Model Variables

Variables	OR	(95% CI)
Satisfaction with care	1.66	(1.08, 2.53)
Language of interview (English vs Spanish)	2.01	(1.06, 3.82)
Income (≤$20,000, >$20,000)	0.57	(0.14, 2.41)*
Self-perceived health status	1.28	(0.74, 2.19)*

*These variables were found to be significant in the logistic regression model but use of adjusted standard errors in calculation of the CI included the null value.

for utilization of folk healers in other ethnicities are not available in the literature.

Mexican Americans may be reluctant to discuss use of curanderos and this could have an impact on the findings of this investigation. However, bilingual interviewers of predominantly Hispanic origin were used in the HHANES. While some reluctance to answer questions regarding curanderismo usage may have occurred, the level of reluctance is thought to be decreased by the cultural and linguistic similarity between the interviewers and the sample persons.

It has been hypothesized[13] that, as a consequence of the use of curanderos, Mexican Americans tend to delay seeking medical care, often until the condition becomes critical. The data presented here do not support that view. If this were true, then differences in time since last health care visit and health status would be evident in the two groups compared. This was not found to be the case.

Regression analyses revealed that adherence to a Mexican cultural orientation (in the form of preference to be interviewed in Spanish) and dissatisfaction with medical care received were related to use of curanderos while ability to pay and availability of care (as indicated by the reporting of a regular source of care) were not so related.

It is possible that those persons who spoke Spanish during the interview are experiencing greater difficulty communicating with health care professionals (predominantly monolingual in English) hence are more dissatisfied with the care they received and seek an alternative source. It is reasonable to assume that those persons who do not have the expendable income to "shop" for a physician would also be more likely to show dissatisfaction with care and seek an alternative source of care, accounting for the marginal predictability of income. Similarly those who feel that they are in poor health may be somewhat more likely to seek persons with whom they can communicate more easily, and whom they perceive as being more empathetic—in this case, curanderos.

This research has been a first step in trying to unravel complex relationships. A major limitation of this research involves the small number of Mexican Americans who reported consulting a curandero; this precluded an examination of possible interactive variations which may have occurred with respect to curandero use. Nevertheless, the data indicate that few Mexican Americans consult folk medicine practitioners, and that those who do so also consult physicians. The research suggests that health care professionals should maintain and strengthen their efforts to deal with linguistic and cultural differences in order to be as effective as science allows.

APPENDIX

Variables Used in the Analyses

Perceived control of future health status was derived from the following question: "How much control do you think you have over your future health: a great deal, some, very little, or none?"

Income was derived from the following question: "Was the total combined family income during the past 12 months more or less than $20,000? Include money from jobs, Social Security retirement income, unemployment payments, public assistance, and so forth. Also include income net from interest, dividend, income from business, farm or rent, and any other money income received."

Determination of insurance coverage was an amalgam of several questions. A positive response to one or more of the following would indicate the presence of coverage, a negative response to all would indicate the absence of insurance coverage:

- "Is sample person now covered by Medicare?"
- "Does sample person have a Medicaid card?"
- "Is sample person now covered by any other public assistance program that pays for health care?"
- "Is sample person now covered by CHAMPUS, which is medical insurance for dependents or survivors of disabled veterans?"
- "Is sample person now covered by any other program that provides health care for military dependents or survivors of military persons?"
- "Is sample person covered under this plan?" (referring to for-profit health insurance plans to which they or their employer may subscribe).

Type of insurance coverage was determined by the presence or absence of preceding items. Not covered by insurance is equivalent to no insurance coverage above.

The *poverty level* variable was a dichotomization of those at or above and those below the poverty index. The poverty index is a ratio of two components: the numerator is the midpoint of the income bracket reported for each family; the denominator is a poverty threshold which varied with the number of persons in the family, the adult/child composition of the family, the age of the reference person, and the month and year in which the family was interviewed.[20]

The *time since last visit* was determined by the compilation of two questions:

- "Now I would like to ask you some questions about your last visit to the (care source as previously determined). How long has it been since that visit? Less than one month; one month but less than six months; six months but less than one year; one year but less than five years; more than five years."

A second question has the same response categories with the addition of the choice of "never", and it reads:

- "Now I would like to ask you some questions about your last visit to any clinic, health center, doctor's office, or other place for health care. How long has it been since your last visit for health care?"

The *source of care variable*, speaking of the "scientific" source of health care, comes from three questions:

- "Is there a particular clinic, health center, doctor's office or other place that you usually go to if you are sick or need advice about your health?"
- Have you visited any other clinic, health center, doctor's office or other place for health care since your last visit to the (regular source)?"
- "How long has it been since your last visit for health care?"

A positive response to the first question and a negative response to the second indicates a regular only source of health care; positive responses to both questions indicates a regular and other source of health care; those who responded positively to the second question but negatively to the first indicate other source only; and, a response of "never" to the third question indicates that the respondent has never visited any source of health care.

The *satisfaction with care variable* is derived from the question:

- "In general, how satisfied were you with the care you received at that visit?" Would you say you were very satisfied, some what satisfied, or not at all satisfied?"

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VI. Alcohol Consumption Patterns among Mexican American Mothers and among Children from Single- and Dual-Headed Households: Findings from HHANES 1982–84

CHRISTINE A. STROUP-BENHAM, MA, FERNANDO M. TREVIÑO, PHD, MPH, AND DOROTHY B. TREVIÑO, PHD

Abstract: Data from the southwestern United States sample of the Hispanic Health and Nutrition Examination Survey were employed to compare the patterns of alcohol use among Mexican American mothers and children in female-headed households with use patterns among mothers and children in couple-headed households. Single female heads of household drank more alcoholic beverages on more days than females from dual-headed households. As a whole, the children of single heads of household still living at home did not demonstrate significantly different drinking patterns from their dual-headed household counterparts. While male children of single-headed households drank more days and total drinks than their dual-headed household counterparts, female children of dual-headed households drank more days and total drinks than female children from single-headed households. [*Am J Public Health* 1990; 80(Suppl):36–41.]

Introduction

The number of single female-headed families in the United States has been increasing during the past three decades. By 1984 over 25 percent of all families that included children were female-headed.[1] Data from Bean and Tienda[2] indicate that 36.5 percent of Puerto Ricans, 18.9 percent of Mexican Americans, and 16.0 percent of Cuban Americans lived in single female-headed families in 1980.

Studies indicate that Hispanic female-headed families are exposed to a variety of potential stressors, such as poverty,[3,4] lack of health insurance,[5] lack of child health care,[6] and lack of social support.[6,7] This suggests they may also be at higher risk for behaviors that have been associated with higher levels of stress, such as alcohol use. Women who had never been married or who were divorced or separated have been found to have lower rates of abstention and higher rates of heavier drinking,[8] especially if cohabitating.

In our analysis of the Hispanic Health and Nutrition Examination Survey data, we attempted to determine if mothers and children in female-headed households (i.e. no male head of household present) displayed alcohol drinking patterns that reflected a greater quantity and frequency of alcohol consumption than did the drinking patterns among the mothers and children from dual-headed households (i.e. both mother and father present) and if factors such as age, acculturation level, education level, and socioeconomic status modified any such relation between household type and drinking patterns. Previous studies[9] have suggested that such factors are partially responsible for differences in behavior observed among single female heads of household and females from dual-headed households.

Methods

The Sample

The Hispanic Health and Nutrition Examination Survey (HHANES) was conducted from 1982 to 1984 by the National Center for Health Statistics (NCHS). The objective of this survey was to secure a representative sample of approximately 12,000 Hispanics in three geographic areas: the southwestern United States, Dade County, Florida, and the New York City metropolitan area.[10] Our research was limited to Mexican Americans (i.e. those with a national origin recode of 1) from the Southwestern sample, since there were fewer cases of single female-headed families among the Cuban and Puerto Rican populations.

The analyses were restricted to those single female-headed (246) and dual-headed (1,022) families in which the mother and at least one of her children under 19 years of age at the time of the interview were sampled (the analytic sample consists of mothers and all of their children who were included as sample people in the HHANES). This strategy assured greater comparability with other investigations of household headship.[11]

Each individual included in the HHANES does not represent the same number of individuals in the actual population, since certain age groups were under- or over-sampled. Therefore, sample weights were used in order to generate correct estimates.[10,12] All analyses were conducted both unweighted and weighted. The unweighted Ns and the weighted statistics are reported unless otherwise stated.

Measures

Individuals ages 12 to 74 were asked questions about alcohol use in the Adult Person Supplement Questionnaire of the HHANES, including:

- on how many of the prior 28 days they had consumed any beer, wine, or liquor;[13]
- total number of drinks (rather than days) summed across all alcohol types consumed during the reference period. One single female head of household (28 years of age, employed, tenth grade education, and living below the poverty level) reported consuming a total of 789 alcoholic drinks during the 28-day reference period;
- the average number of "drinks per occasion" (computed, by dividing the total number of alcoholic drinks consumed during the reference period by the total number of days the subject drank).

Those who had drunk any type of alcohol during the reference period were classified as "drinkers," all others were "abstainers." In females, "high drinking" was defined as having an average of four or more alcoholic drinks per occasion during the reference period and "low drinking" was defined as having one to three drinks per occasion. In males, the corresponding definitions were six or more and one to five. These created variables parallel those employed by Markides, *et al*, in their 1988 three-generation study.[14]

NOTE: Author affiliations are listed elsewhere under CONTRIBUTORS.

The HHANES included a modified version of an acculturation scale developed by Cuellar, *et al*,[15] which has been extensively utilized among Mexican American populations. The modified scale is comprised of eight items pertaining to language preference, language spoken, language subject reads better, language subject writes better, ethnic identification of the subject, ethnic identification of the subject's mother, ethnic identification of the subject's father, and one item that combined the place of birth of the subject's parents. This scale has a minimum value of 1 and a maximum value of 4.9, with higher scores indicating higher levels of acculturation.

Age refers to the actual age of the individual at time of examination. Education is the number of years of school completed. Employment status was coded 1 for employed (part- or full-time) and 0 for non-employed. Poverty was coded as 1 for persons living below the federal poverty cut point and as 0 for those at or above poverty level. Income reflects the combined family income (categorized) over the 12 months prior to the interview. All of the dual heads of household were self-identified as a spouse of a head of household or as head of household with a spouse.

None of these variables had a non-response rate of greater than 10 percent for the mothers and 12 percent for the children. The majority of the questions were answered by nearly all subjects in the sample. Therefore, the effects of non-response are believed to be minimal.

Analyses

The variables (total drinks, total number of drinking days, and drinks per occasion) were examined within each household type for three age categories among the mothers, 17–29, 30–39, 40–74 (approximate tertiles), using analysis of variance.

Analysis of covariance was employed to test the hypothesis that mothers and children of single-headed households would consume more alcohol than those in dual-headed households when controlling for age, acculturation, combined family income, and education. Multiple classification covariance analysis (see, e.g., Nie, *et al.*[16]) produces adjusted means across the values of each independent variable. Analysis of covariance (ANCOVA) employed total number of drinks as the dependent variable with dual vs single household type as the independent variable. The covariates—including combined family income, acculturation level, education level, and age—were then entered into the procedure. Adjusted means were obtained for each level of the independent variable. This same procedure was employed for two additional ANCOVAs using, respectively, the dependent variables of total number of days drank and number of drinks per occasion.

Analysis of covariance was also employed to explore children's drinking patterns. Three separate analyses were conducted, using total number of drinks, total number of days, and number of drinks per occasion, respectively, as the dependent variables and household type as the independent variable. The covariates in each analysis were age, education, gender, combined family income, and acculturation.

Regression coefficients and their standard errors are presented for the covariates of each analysis of covariance. Standard errors have been adjusted for design effects[10,12] (design effects were calculated employing the Taylor approximation technique available with the SESUDAAN program[17]).

TABLE 1—Sociodemographic Variables of Females and Children (ages 12–18 years) of Single and Dual Households

	Mothers		Children	
Variables	Single Household (n = 246)	Dual Household (n = 1022)	Single Household (n = 201)	Dual Household (n = 624)
Age Categories				
% 12–14 years	NA	NA	44.2	47.0
% 15–16 years	NA	NA	29.1	28.0
% 17–29 years	26.3	36.4	26.7	25.0
% 30–39 years	35.3	37.7	NA	NA
% 40–74 years	38.4	25.9	NA	NA
Education Level				
Mean	8.6	9.1	8.9	8.7
Standard Error	.24	.11	.13	.08
% In Poverty	64.5	26.3	70.3	31.2
Acculturation				
Mean	2.6	2.5	3.3	3.2
Standard Error	.07	.03	.05	.03
% Employed	58.2	44.3	NA	NA
% Males	NA	NA	45.7	51.7
Family Size				
Mean	3.7	4.8	4.9	5.9
Standard Error	.12	.05	.18	.07

SOURCE: NCHS Hispanic Health and Nutrition Examination Survey

Results

An examination of marital status indicated that the majority of single female heads of household were divorced (34.5 percent) or separated (33.1 percent). The remainder of the single females were either widowed (18.1 percent) or had never been married (14.3 percent). Examination of the marital status of the children of these women revealed that just nine were married and four were separated while the remainder had never been married. Tables 1 and 2 present weighted means and percents for the alcohol variables as well as sociodemographic characteristics of mothers and children controlling for household type. More single mothers (and their children) were living in poverty and were more likely to be employed in comparison with their dual-headed counterparts (Table 1). The observed differences between mothers and children regarding percent impoverished and family size are a function of a lack of a one-to-one correspondence between mothers and children (The statistics would be the

TABLE 2—Alcohol Consumption Variables of Females and Children (ages 12–18 years) of Single and Dual Households

	Mothers		Children	
Variables	Single Household (n = 246)	Dual Household (n = 1022)	Single Household (n = 201)	Dual Household (n = 624)
Total No. Days Drank				
Mean	2.0	.9	.7	.7
Standard Error	.32	.11	.19	.11
No. Drinks per Occasion				
Mean	4.7	2.7	4.1	3.8
Standard Error	.47	.13	.54	.38
Total No. Drinks				
Mean	12.4	2.5	3.3	2.5
Standard Error	3.90	.30	1.10	.46
% High Drinkers	46.0	23.8	38.4	23.4
Drinker vs Abstainer				
% Drinkers	44.4	25.0	17.6	17.0

SOURCE: NCHS Hispanic Health and Nutrition Examination Survey same if there was a one-to-one correspondence.)

There were 92 single female heads and 214 females of couple-headed households who were drinkers; 29 children who were drinkers were in single households and 84 were in dual-headed households; 46 percent of the single female heads of household who were drinkers were high drinkers compared to 24 percent of the female drinkers in dual-headed households (Table 2). Examination of the percentages of drinkers across marital status of single female heads of household indicated that there were a greater proportion of drinkers among those who were divorced (47.9 percent) or separated (41.0 percent) than those who were never married (33.3 percent) or were widowed (32.9 percent).

When sociodemographic characteristics of mothers from single and two-parent households were examined controlling for drinker type (i.e. abstainer vs drinker), families of drinking women tended to be smaller than families of abstaining women. Women who were drinkers had higher education levels and were less likely to be impoverished than abstainers. Drinkers from both household types had a higher level of acculturation and a greater percentage employed than abstainers. The analysis of variance (ANOVA) procedures indicated that there were no significant differences between the three age categories of mothers in single- and dual-headed households for the dependent variables of total number of days drank, total number of drinks, and drinks per occasion.

The proportion of children living in poverty was less among drinkers than abstainers. The children who were drinkers were slightly older. In order to determine if the proportion of drinkers varied across age groups among the children, the percentage of drinkers in three age categories (i.e. 12–14, 15–16, 17–19 years) was computed. While cell sizes were small for the youngest age categories, no significant differences in the percent of drinkers across age groups were found, even when controlling for household type.

Results of the ANCOVA procedure indicated that the weighted crude mean total drinks for single heads of household was 13.17 and for females of dual-headed families it was 2.46. When the covariates—highest level of education achieved, combined household income, age, and acculturation—were taken into consideration, the resulting weighted adjusted means were 21.91 drinks for singles and .32 drinks for duals (Table 3). The second panel of Table 3 provides the regression coefficients (bs) and standard errors (adjusted for design effects) of the covariates. These results indicate that household type had a strong negative association with total number of drinks consumed, meaning that single heads of household drank more than women from couple-headed families. Acculturation was found to have a positive relationship with total number of drinks, indicating that women who were more acculturated tended to drink more.

Single female heads of household drank a weighted mean of 2.12 days and females of couple-headed households drank a weighted mean of .92 days; adjusted weighted means were 3.34 and .62 days, respectively (Table 3). The strongest relationship found between covariate and number of days drank was a negative one with household type. The remaining three covariates had weaker positive associations with the dependent variable.

The adjusted weighted means of number of drinks among females per occasion were 6.06 for single household heads and 2.23 for dual household heads (Table 3). As with the previous analyses, household type was discovered to have a strong inverse relationship with the dependent variable, number of drinks per occasion.

Because one single female head of household reported 789 drinks for the 28-day reference period, we analyzed the data after excluding her. In addition, a logarithmic transformation of the three dependent variables was computed from the data (including this single head of household) because of the skewed distribution. These new values were utilized in the three ANCOVAs. These procedures produced results comparable to those shown in Table 3.

Since acculturation appeared to have the most influence (other than household type) with respect to the dependent variables, further analyses were undertaken to uncover its relationship with these measures. It was found that among both single and dual female heads of household, acculturation levels were significantly higher among those who consumed more alcohol.

Analysis of covariance was also utilized to examine the children's (ages 12 to 18 years) continuous drinking measures. Covariates included age, education, gender, combined

TABLE 3—Analysis of Covariance of Mothers' Three Continuous Drinking Measures

	Female Heads of Household								
	Total Number Drinks			Total Number Days Drank			Drinks per Occasion		
	N	Crude	Adjusted*	N	Crude	Adjusted*	N	Crude	Adjusted*
Household Type									
Single	202	13.17	21.91	202	2.12	3.34	84	4.83	6.06
Dual	827	2.46	.32	825	.92	.62	195	2.76	2.23
Total (Grand mean)		4.57			1.16			3.38	
Multipe R^2		.028			.057			.139	
		b	**SE**		**b**	**SE**		**b**	**SE**
Covariates									
Age		−0.107	0.091		−0.004	0.011		−0.025	0.025
Education		0.031	0.311		0.025	0.038		−0.108	0.119
Income		−0.016	0.107		0.045	0.037		−0.036	0.021
Acculturation		1.169	1.092		0.419	0.148		0.259	0.260
Household Type		−10.881	2.334		−1.519	0.340		−1.765	0.564

*Adjusted for acculturation, age, education, and combined family income.
SOURCE: NCHS Hispanic Health and Nutrition Examination Survey.

family income, and acculturation level, with household type as the independent variable. The weighted adjusted mean number of total drinks during the reference period was 3.80 for children in single-headed households and 2.42 for children of couple-headed households, .75 and .68 respectively for total drinking days, 3.59 and 4.17 respectively for drinks per occasion (Table 4).

None of these ANCOVAs examining the continuous drinking variables for the children indicated that household type had a strong relationship with the dependent variables. However, all three analyses exploring the children's drinking patterns demonstrated strong associations between the covariates age, sex, and acculturation and the dependent variables (Table 4). In order to discover how these covariates were related to the continuous drinking variables, several analyses were done. Male children of single heads of household were found to drink more and on more occasions than male children from two-parent homes (Table 5). Whereas, female children of dual-headed households drank more than those from single households. No significant differences were found between high vs low drinkers or abstainers vs drinkers with respect to level of acculturation or age. Drinking children were slightly older than non-drinking children in both single- and dual-headed households (16.4 years compared to 14.5 years).

In order to determine if drinking status among children was related to drinking status of mother, proportions of drinking children (controlling for gender) were computed for abstaining and drinking mothers of each household type. As shown in Table 6, percentages of drinking daughters were similar among abstaining mothers from couple-headed households (11.9 percent) and abstaining single heads of household (9.2 percent). A greater proportion of drinking daughters came from two-parent homes with a drinking mother (23.3 percent) than came from single-parent homes with a drinking mother (11.5 percent). Among single-headed families with abstaining mothers, 35.6 percent of the sons were found to drink, while 16.9 percent of the sons from dual-parent homes with abstaining mothers did so. Within single-parent families having a drinking mother, 20.7 percent of the sons drank. It was found that 26.2 percent of the sons in two-parent homes with a drinking mother were drinkers also. Analyses indicate that these gender differences do not appear to be related to a variation in age distribution.

Discussion

As hypothesized and as found in other ethnic groups, there was greater use of alcohol among parents in non-intact families compared to those from intact families,[18,19] even when age, acculturation level, education, and income were taken into consideration. Acculturation did appear to play a role in drinking behaviors, since women drinkers of both single- and dual-headed households had higher levels of acculturation.

It has been suggested in other studies that more acculturated Mexican American women do have a tendency to have higher alcohol consumption.[13,14,20,21] Higher levels of acculturation do not explain why single female heads of household consume much more alcohol than their dual counterparts, however. Stressors of being a single female head of household need to be explored to ascertain what, if any, role they have in exacerbating drinking behaviors of single female heads of household.

It has been suggested by others[9] that when socioeconomic status is controlled, differences between single- and dual-headed households become non-significant. However, this study does not support such findings with respect to alcohol consumption patterns. In fact, we found that when such factors as income and education are controlled for, the differences between women of the two household types became more pronounced. Therefore, other elements must be contributing to the heavier alcohol consumption among single female heads of household. This conclusion is supported by the relatively small values of the adjusted R^2's for the ANCOVAs—suggesting that the independent variables explain little of the variance.

Since we have no data on initiation of alcohol use in conjunction with changes in marital status, it may be that marriage breakdown was due partially to the drinking behav-

TABLE 4—Analysis of Covariance of Children's (ages 12–18) Three Continuous Drinking Measures

	Children (ages 12–18)								
	Total Number Drinks			Total Number Days Drank			Drinks per Occasion		
	N	Crude	Adjusted*	N	Crude	Adjusted*	N	Crude	Adjusted*
Household Type									
Single	164	3.39	3.80	163	.69	.75	26**	4.02	3.59
Dual	490	2.56	2.42	490	.70	.68	70	4.01	4.17
Total (Grand mean)		2.77			.70			4.01	
Multiple R^2		.086			.105			.220	
		b	**SE**		**b**	**SE**		**b**	**SE**
Covariates									
Age		1.627	0.642		0.409	0.120		0.754	0.493
Education		−0.133	0.604		−0.071	0.111		−0.213	0.519
Gender		−3.061	1.147		−0.643	0.224		−1.856	0.735
Income		−0.009	0.060		0.018	0.015		−0.043	0.042
Acculturation		1.503	0.677		0.304	0.148		0.952	0.530
Household Type		−0.551	1.428		−0.076	0.333		0.591	0.978

*Adjusted for acculturation, age, sex, education, and combined family income.
**Statistic does not meet a minimum reliability standard (due to small cell size).
SOURCE: NCHS Hispanic Health and Nurtrition Examination Survey

TABLE 5—Comparison of Continuous Drinking Measures for Children (ages 12–18) by Household Type

	Male Children				Female Children			
	Single Household		Dual Household		Single Household		Dual Household	
	N	Weighted Mean	N	Weighted Mean	N	Weighted Mean	N	Weighted Mean
Total Drinks	80	6.30	270	3.50	100	.75	275	1.48
S.E.		2.34		.82		.37		.37
Total Days Drank	80	1.18	270	.94	99	.24	274	.44
S.E.		.40		.19		.10		.09
Drinks per Occasion	20**	4.66	41**	4.73	8**	2.68	38**	2.67
S.E.		.65		.64		.71		.29

**Statistic does not meet a minimum reliability standard (due to small cell size).
SOURCE: NCHS Hispanic Health and Nutrition Examination Survey

TABLE 6—Weighted Percentages of Children Classified as Drinkers Controlling for Drinking Status of Mother

	Children Classified as Drinkers	
Status of Mothers	% Male	% Female
Single Household		
Abstainer	35.6	9.2
Drinker	20.7	11.5
Dual Household		
Abstainer	16.9	11.9
Drinker	26.2	23.3

SOURCE: NCHS Hispanic Health and Nutrition Examination Survey

iors of these females or that being single allowed these women more opportunity to consume alcohol.

The findings regarding female children are surprising. One would expect that daughters of single female heads of household would drink more than daughters of dual household heads, but we found the reverse to be the case. Additionally, sons of abstaining single mothers were more likely to be drinkers in comparison with both sons of abstaining dual mothers and sons of drinking mothers of either household type. Of course, within couple-headed families, children have a father who may act as a role model for their drinking behaviors. This could explain the findings regarding children if fathers were more likely than mothers to be drinkers.

In summary, these data suggest that single female heads of household drink significantly more than women in couple-headed families. This research indicates that single female heads of household are at increased risk for greater consumption of alcohol. While a lower percentage of mothers from couple-headed households consume alcohol, their numbers are greater. This information may be of assistance to those who implement preventive and intervention measures and treatment, especially since data indicate that alcohol-related mortality is particularly high among Mexican Americans.[22] Additionally, further research needs to be done to investigate the factors contributing to the drinking behaviors among these women, such as change in marital status and psychological stressors. In this way, treatment and intervention may become more effective.

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VII. Acculturation and Alcohol Consumption in the Mexican American Population of the Southwestern United States: Findings from HHANES 1982–84

Kyriakos S. Markides, PhD, Laura A. Ray, MPA, Christine A. Stroup-Benham, MA,
and Fernando Treviño, PhD, MPH

Abstract: Data from the Southwestern sample of the Hispanic HANES are employed to evaluate the relationship of acculturation into the larger society with alcohol consumption. As in previous work, acculturation was not found to be related to alcohol consumption of Mexican American men, but was positively related to the consumption of younger Mexican American women. Among middle-aged women, acculturation was not important. However, we found evidence that middle-aged women might be turning to alcohol in response to marital disruption and poverty. In addition, middle-aged women who are not employed are less frequent drinkers but those who drink are heavier drinkers than employed women. [*Am J Public Health* 1990; 80(Suppl):42–46.]

Introduction

In recent years, a number of papers focusing on predictors of alcohol consumption among Mexican Americans have been published. By far, the leading predictor of interest has been "acculturation" into the larger society.[1–4] The concept of acculturation has actually assumed great importance in studies of Mexican Americans in general during the 1980s, particularly in research on mental health.[5,6] Despite its increasing importance, little progress has been made in both conceptualizing and measuring acculturation.

Gilbert and Cervantes,[2] who have reviewed much of the literature on acculturation and alcohol consumption among Mexican Americans, have suggested two models generally followed.

- The *acculturative stress model,* according to which higher and/or more frequent alcohol consumption is the result of the many stresses accompanying the process of acculturation into the larger society. Higher drinking would thus be expected in persons at the middle of the acculturation continuum, and not at the extreme ends.
- The *acculturation model*, on the other hand, predicts that alcohol consumption patterns of Mexican Americans would reflect the extent to which they have adopted the drinking norms and practices of the larger society.

The literature has been mixed with regard to support of these perspectives.[1–2,4] In general, however, it appears that the "acculturation" model tends to be more predictive of the behavior of women.[1,2] With regard to men, the literature has been less clear-cut.[1]

Working from these two perspectives, Markides, *et al*,[3] hypothesized that the "acculturative stress" model would have relevance to the drinking experience of men, while the "acculturation" model would have more relevance for women. The latter hypothesis was based on the relatively low drinking rates of Mexican American women. Since, as a group, they drink little, women would be less likely to turn to alcohol as a way to relieve any stress associated with acculturation.

Markides, *et al*,[3] tested these hypotheses using data from a three-generation study of Mexican Americans living in the San Antonio area. Among men, they found that acculturation was related to lower alcohol consumption in the middle generation (middle-aged people). Since virtually all of the middle-generation subjects were born in the United States, people at the low end of the acculturation scale were at the middle of the acculturation continuum, and thus, the results were interpreted as consistent with the "acculturative stress" hypothesis. Among women, acculturation was positively related with alcohol consumption in the younger generation, as predicted by the acculturation hypothesis. However, middle-generation women behaved much like their male counterparts as they appeared to drink more if they were less "acculturated."

This paper attempts to replicate the above study using data from the more representative southwestern United States sample from the Hispanic Health and Nutrition Examination Survey (HHANES). Thus, the analyses are conducted separately by sex and for three age groups: 20–39, 40–64, and 65–74 years. These correspond roughly with the ages of the three-generations in Markides, *et al*.[3] However, the middle and younger age groups have more variation on acculturation since they also include first-generation (foreign-born) Mexican Americans.

Methods

The data were collected as part of the Adult Sample Person Supplement of the southwestern sample of the Hispanic HANES.[5] Our analysis is limited to the Mexican American respondents from the five Southwestern States (Texas, New Mexico, Colorado, Arizona, and California). The Adult Sample Person Supplement Questionnaire contains a long series of questions regarding alcohol consumption. We used several variables which enabled us to construct similar measures to those employed by Markides, *et al*.[3] Respondents were asked how many days altogether they drank beer during the four-week period ending on the Sunday prior to the interview. Next, they were asked how many beers they drank per day during the days they drank beer. They were also asked how many beers they drank in total during the four-week period. These questions were repeated with regard to wine and with regard to "liquor such as: whiskey, rum, gin, vodka or tequila." Respondents were also asked on how many days they drank any kind of liquor.

Using the above information, we were able to create the following variables: *Frequency* refers to the number of days a person had consumed any alcoholic beverages during the 28-day reference period. *Volume* refers to the average num-

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NOTE: Author affiliations are listed elsewhere under CONTRIBUTORS.

ber of drinks consumed per day by those who drank. *Total Drinks* refers to the total number of drinks consumed during the 28-day period by all respondents. *High Drinking* refers to the percentage of men who drank six or more drinks a day when they drank, and women who drank four or more drinks a day when they drank. *Drinkers vs Abstainers* refers to the percent of respondents who had drunk any alcoholic beverages during the 28-day reference period (coded 1 for drinkers and 0 for abstainers).

The key independent variable in our analysis is "acculturation." The Hispanic HANES utilized a shortened eight-item version of the Cuellar acculturation scale, which is the most widely used scale in research with Mexican Americans.[7] Items included captured the respondents' language use and preference, ethnic identification and nativity, both for the respondent and his/her parents. Internal consistency reliability for the scale is in excess of .90.

Other variables employed were more straightforward. *Age* was actual age at the time of the interview; *Education* refers to years of school completed; *Marital Status* was coded 1 for married persons (spouse present) and 0 for all others; *Poverty* was coded 1 for people living below the federal poverty line and 0 for all others; finally, *Employment* was coded 1 for persons who were employed (full-time or part-time) at the time of the interview and 0 for all others.

The five alcohol measures discussed earlier were regressed on the acculturation scale and the five control variables: age, education, marital status, poverty, and employment. The analysis was performed separately by sex and for three age groups: ages 20-39, 40-64, and 65 and over. As mentioned above, these age groups correspond roughly to the three generations in Markides, *et al*,[3] who have shown that the influence of acculturation on alcohol consumption varies by generation (age) and sex. All analyses were performed with appropriate adjustments for sample weights and design effects. Design effects were necessary because the Hispanic HANES employed a complex multi-stage probability sample design. Sample weights were normalized by multiplying the number of subjects in each age/gender group employed in the analysis with each respondent's weight divided by the sum of the weights for the group. Design effects were computed using the Taylor Approximation of the SESUDAAN program and were subsequently employed to adjust the variances in the regression analyses. Examination of correlation and covariance matrices did not reveal evidence of multicollinearity (Intercorrelation tables are not presented but are available from the authors upon request). Regression analyses employed listwise deletion of missing values which led to the loss of cases ranging from 9 to 15 percent.

The analysis is aimed at evaluating the usefulness of the "acculturative stress" and "acculturation" perspectives.[2] As mentioned earlier, we expect the former to have more relevance for men and the latter to have more relevance for women. However, based on our unexpected findings with middle-aged women in the three-generation study,[3] we also examine the possibility that some middle-aged women in this sample might resort to alcohol consumption as a way to relieve stress related to acculturation or other factors.

Results

Table 1 presents means and percents on the five alcohol measures described above by sex and age group. Maximum sample sizes for each group are shown at the top. These ranged from 1,016 younger females to 84 older males. Number of cases for the alcohol variables vary because of missing values and because nondrinkers are not included in the measurement of some variables. For example, volume is the average number of drinks per day when a person drank. Only people who had drunk during the 28 days prior to the interview are included. This means considerably small sample sizes, particularly among women and older people, most of whom are not drinkers.

As Table 1 shows, drinking is considerably more frequent and heavier among younger and middle-aged men as previous studies have shown. Although these men do not appear to be very frequent drinkers when compared with men in the general population, they drink heavily when they drink, which is consistent with findings from previous studies.[2,4,8]

TABLE 1—Alcohol Consumption Variables and Sociodemographic Variables by Gender and Age Group

	Ages 20–39		Ages 40–64		Ages 65–74	
Variables	Male (N = 822)	Female (N = 1016)	Male (N = 558)	Female (N = 726)	Male (N = 84)	Female (N = 119)
Frequency						
(mean Number of Days)	6.4	1.4	7.8	1.4	4.0	.92
(N)	(821)	(1015)	(556)	(725)	(84)	(119)
Volume						
(Mean)	4.9	3.2	4.4	3.0	3.4	1.9
(N)	(587)	(331)	(368)	(182)	(38)	(17)
Total Drinks						
(Mean)	28.0	4.8	33.9	5.6	13.1	1.5
(N)	(821)	(1016)	(557)	(726)	(84)	(119)
High Drinking						
(% 6+ Males, 4+ Females)	31.4	28.1	24.6	24.9	20.7	0
(N)	(587)	(331)	(368)	(182)	(38)	(17)
Drinkers vs. Abstainers						
(Percent drinkers)	72.6	34.5	69.5	26.3	46.8	14.2
(N)	(822)	(1016)	(558)	(726)	(84)	(119)
Age	28.5	28.6	50.2	50.5	69.1	69.8
Education (mean yrs)	10.4	10.2	8.2	7.6	5.4	4.8
Percent Married	70.9	67.2	89.4	71.6	81.5	42.6
Percent in Poverty	22.0	30.0	21.1	31.5	44.1	58.1
Acculturation	2.6	2.7	2.4	2.3	2.2	2.2
Percent Employed	94.2	53.5	84.1	45.9	13.5	6.8

Table 1 also presents data (means or percents) on several sociodemographic variables included in the regression analyses. As would be expected, acculturation is highest in the younger age group and lowest in the oldest. Also, the poverty rate is higher among women and the elderly. Other variables are distributed by sex and age as might be expected. As expected, consumption is much lower in the older age group, and much higher among men than among women in all three age groups.

Table 2 presents the results of the regression analyses for men and Table 3 of the analyses for women. The analyses were conducted separately for three age groups. The top panel of Table 2 shows the results for younger men (ages 20–39). As will be seen, acculturation is not associated with any of the alcohol variables in younger men. Of the other variables, marital status is negatively related to frequency and total drinks consumed (married men drink less); education is negatively related to volume of consumption and the probability of being a high drinker; similarly, living in poverty is positively associated with volume of consumption and probability of being a high drinker; finally, employment is associated with greater frequency of drinking.

The results for middle-aged men (Table 2) show no notable associations between acculturation and any of the alcohol measures. Age was negatively associated with the probability of being a drinker, while education was positively related with the same variable and with frequency of consumption. On the other hand, education was negatively associated with volume of consumption and the probability of being a high drinker. Married men had lower volume of consumption and consumed fewer total drinks and poverty was associated with less frequent consumption.

As would be expected, fewer associations were observed with older men, although this is partly the result of small numbers. Table 2 shows a negative association between age and the probability of being a drinker while marital status is positively associated with the same variable.

Table 3 shows the results for women. Among younger women, acculturation is positively associated with frequency of consumption and the probability of being a drinker. More educated younger women are more likely to be drinkers but they consume a lower volume of alcohol and are less likely to be high drinkers. Marital status is negatively associated with all five alcohol variables. Finally, employment is positively related to the probability of being a drinker.

Among middle-aged women, acculturation is positively related to the probability of being a drinker, but not to any of the other alcohol variables. Age was negatively associated with the probability of being a drinker, volume of consumption, total drinks consumed, and the probability of being a high drinker. Marital status was negatively associated with volume of consumption, the probability of being a high drinker, and the probability of being a drinker. Poverty was negatively associated with the probability of being a drinker; however, it was positively associated with volume of consumption and the probability of being a high drinker. Finally, employed middle-aged women were more likely to be drinkers and to be frequent drinkers. However, they consumed a lower volume of alcohol and were less likely to be high drinkers.

Among older women we find that education is positively associated with the probability of being a drinker. No other associations are noted but, again, the small number of cases is a problem.

Discussion

As shown in a number of previous studies,[1,2] acculturation is not related to alcohol consumption among men in any consistent manner. We expected a negative association between acculturation and alcohol consumption among men,

TABLE 2—Results of Multiple Regression Analysis of Alcohol Consumption: Males

	Frequency		Volume		Total Drinks		High Drinking		Drinkers vs Abstainers	
	b	SE	b	SE	b	SE	b	SE	b	SE
					Ages 20–30					
Age	.149	.082	.018	.049	.264	.430	.017	.025	.013	.018
Education	.009	.140	−.371	.072	−1.051	.712	−.134	.039	.048	.033
Marital Status	−1.931	.860	−.958	.554	−11.611	4.953	.203	.289	−.228	.214
Poverty	−.360	1.043	1.258	.681	5.470	5.132	.583	.350	−.183	.203
Acculturation	.663	.439	.394	.271	4.040	2.574	.171	.138	.057	.103
Employment	3.602	.512	−.101	1.092	14.864	8.687	.192	.565	.183	.448
(N)	(746)		(544)		(746)		(544)		(747)	
					Ages 40–64					
Age	−.058	.075	−.036	.033	−.781	.469	−.027	.025	−.04[illegible]	.016
Education	.342	.126	−.204	.060	1.420	.839	−.144	.042	.100	.030
Marital Status	−2.553	1.840	−2.152	.692	−32.179	10.223	−.148	.456	−.272	.381
Poverty	−2.760	1.157	.836	.589	4.419	8.002	.451	.384	−.386	.250
Acculturation	−.870	.764	.399	.283	−.058	4.796	.202	.199	−.117	.133
Employment	−.881	1.748	−.698	.879	−1.508	9.804	−.249	.530	.590	.308
(N)	(491)		(326)		(491)		(326)		(493)	
					Ages 65–74					
Age	−.016	.386	−.110	.189	−.008	1.739	−.562	.336	−.216	.108
Education	.178	.265	−.166	.187	.164	1.205	−.041	.309	.075	.080
Marital Status	3.766	2.173	.963	1.566	10.750	11.167			1.474	.742
Poverty	−.167	1.867	−.520	1.174	1.938	8.929	−.886	1.713	.229	.714
Acculturation	−.870	1.189	.765	.711	6.781	5.401	.982	1.149	.081	.361
Employment	.881	2.739	−.174	1.798	−9.401	12.440			−.607	.859
(N)	(73)		(32)		(73)		(32)		(73)	

NOTE: OLS regression coefficients are presented for frequency, volume and total drinks, and logistic regression coefficients for high drinking and drinkers vs abstainers.
$^{*}p < .05$, $^{**}p < .01$

TABLE 3—Results of Multiple Regression Analysis of Alcohol Consumption: Females

	Frequency		Volume		Total Drinks		High Drinking		Drinkers vs Abstainers	
	b	SE	b	SE	b	SE	b	SE	b	SE
Ages 20–30										
Age	.026	.031	.010	.034	.105	.207	−.003	.036	.018	.021
Education	.074	.051	−.184	.061	−.097	.432	−.120	.063	.081	.036
Marital Status	−1.221	.318	−.947	.426	−7.036	2.058	−.621	.427	−.964	.209
Poverty	−.442	.348	.869	.516	1.999	3.133	.571	.439	−.242	.253
Acculturation	.433	.158	.416	.217	1.591	1.137	.171	.186	.431	.102
Employment	.457	.316	−.175	.423	2.256	2.062	−.359	.433	.462	.227
(N)	(919)		(305)		(919)		(305)		(920)	
Ages 40–64										
Age	−.026	.028	−.136	.045	−.580	.279	−.084	.035	−.037	.016
Education	.063	.052	−.043	.093	.081	.528	−.045	.071	.056	.030
Martial Status	−.326	.425	−1.394	.706	−8.425	3.888	−1.468	.459	−.796	.240
Poverty	−.109	419	2.384	.823	7.634	4.416	1.468	.479	−.551	.200
Acculturation	.422	.229	.256	.404	4.004	2.296	−.130	.271	.343	.127
Employment	.921	.385	−1.413	.737	3.683	3.676	−1.263	.468	.893	.248
(N)	(621)		(158)		(621)		(158)		(622)	
Ages 65–74										
Age	.249	.158			.246	.242			.019	.100
Education	.131	.135			.174	.207			.303	.141
Marital Status	−.111	.943			−.584	1.439			−.955	.696
Poverty	−.238	.964			.410	1.514			−.414	.662
Acculturation	.620	.554			1.168	.845			.485	.376
Employment	2.314	1.888			2.912	2.965			.235	1.177
(N)	(102)				(102)				(102)	

NOTE: OLS regression coefficients are presented for frequency, volume and total drinks, and logistic regression coefficients for high drinking and drinkers vs. abstainers.
*$p < .05$, **$p < .01$

particularly in middle-age, as [illegible]he case in the three-generation study by Markides [illegible]l.[3] The lack of such negative association may very we[illegible] result from the higher variance on the acculturation scale among middle-aged men in this study than the second generation men in the three-generations study, virtually all of whom were born in the United States.

With regard to women, acculturation appears to be more predictive, as shown by previous research.[1–3] As found by Markides, *et al*,[3] a positive association is observed in younger women with both frequency and volume of consumption, as well as the probability of being a drinker. However, no association is observed among middle-aged women. Furthermore, no tendency is observed toward a negative association between acculturation and alcohol consumption in middle-aged women as was the case with middle-generation women in Markides, *et al*.[3] Although this difference might be again the result of greater variance in acculturation in the present study, there is no evidence that middle-aged Mexican American women are turning to alcohol as a way to relieve "acculturative stress." However, there is evidence that middle-aged women might be turning to alcohol as a way to relieve stress resulting from other factors, such as poverty and marital disruption. Thus, although poverty is negatively associated with the probability of being a drinker, among middle-aged women who drink, poverty is positively related to the volume of consumption, total number of drinks, and the probability of being a high drinker. Similarly, non-married middle-aged women (mostly divorced or widowed) were heavier drinkers. Finally, the effect of employment is in a direction suggesting heavier drinking among housewives, who are presumably more traditional. Even though middle-aged employed women were more likely to be drinkers and were more frequent drinkers, of those who were drinkers, housewives drank more heavily.

In sum, there is only partial support for an "acculturation" model with regard to alcohol consumption for women and no consistent support for an "acculturative stress" model for either sex. We hope to have demonstrated that the influence of acculturation on alcohol consumption among Mexican Americans is highly complex and depends on such factors as variation in the acculturation continuum, age, sex, and type of measure of alcohol consumption.

We believe that our findings also have implications for the broader literature on the health and mental health of Hispanics where a lively interest has been shown in the concept of acculturation. This research has yielded unclear results with regard to the role of "acculturative stress," because of conceptual and measurement problems, and difficulties in establishing exactly at which point on the acculturation continuum stress is at its highest.[7]

We also feel that studies of the health behaviors and health status of Mexican Americans and other ethnic populations will benefit from introducing a life course perspective into their approach to the data analysis. For example, there is evidence that the concept of acculturation has different meaning and influence at different stages of the life course,[10] as well as for the two sexes, as the present analysis has shown.

ACKNOWLEDGMENTS

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VIII. Patterns of Cigarette Smoking among Hispanics in the United States: Results from HHANES 1982–84

Suzanne G. Haynes, PhD, Clair Harvey, Henry Montes, Herbert Nickens, MD, and Bernice H. Cohen, PhD

Abstract: In the 1982–84 Hispanic Health and Nutrition Examination Survey, the prevalence of cigarette smoking was examined among Mexican Americans, Puerto Ricans, and Cuban Americans in the United States. Among 20–74 years olds, the age-adjusted smoking rates for Mexican American, Puerto Rican, and Cuban American men were high—42.5, 39.8, and 41.6 percent, respectively. Quite striking among Cuban American men was the high smoking rate among 20–34 year olds (50.1 percent), the highest smoking rate in the three Hispanic groups compared. The age-adjusted smoking rates for Mexican American, Puerto Rican, and Cuban American women were much lower than those for men—23.8, 30.3, and 24.4 percent, respectively. Both Puerto Rican and Cuban American men were more likely to be heavy smokers (52.3 and 64.1 percent, respectively, smoking a pack or more a day) as compared to the Mexican Americans (33.8 percent smoking a pack or more a day). The pattern was the same for women, with Mexican American women being lighter smokers (18.8 percent smoking a pack or more a day) as compared to heavy smoking among Puerto Rican and Cuban American women (35.1 and 48.6 percent, respectively, smoking a pack or more a day). Given the health hazards of smoking, future research and intervention are required for those groups with high exposure to cigarette smoking. [*Am J Public Health* 1990; 80(Suppl):47–54.]

Introduction

In recent years, several studies have shown that the prevalence rates of smoking among Hispanics in the United States are relatively low in comparison to non-Hispanic Whites and Blacks. Most studies have examined Hispanic populations either in specific geographic areas or in particular ethnic groups, such as Mexican Americans in Sacramento, California[1–4] or in San Antonio, Texas.* National data have come primarily from the National Health Interview Survey (NHIS) which uses the term "Spanish Origin" to describe those of Hispanic ancestry.[5] These data may not reflect the true rate of smoking in the Hispanic US population since the NHIS data have not been collected with a Spanish language instrument nor have the interviewers been bilingual. In general, the number of Hispanics in most studies has not been large enough to be representative or to produce reliable estimates for analysis. This is particularly relevant if one attempts to disaggregate Hispanic data among various population subgroups, e.g., Mexican Americans, Puerto Ricans, and Cuban Americans.

The Hispanic Health and Nutrition Examination Survey (HHANES) was a probability-based survey of three distinct Hispanic populations in the United States.[6] For HHANES, the targeted population consisted of three groups of civilian noninstitutionalized persons ages six months to 74 years, from three areas of the US. The three groups were: 1) Mexican Americans residing in selected areas of Texas, California, Colorado, New Mexico, and Arizona (otherwise known as the Southwest); 2) Cuban Americans residing in Dade County (Miami), Florida; and 3) Puerto Ricans residing in the greater New York City area.

Smoking among Hispanics is currently gaining greater recognition as a public problem since cigarette manufacturers are now aggressively promoting cigarettes to the Hispanic population.[7,8] A recent report noted that these promotional efforts include advertisements in publications aimed at minority groups, sponsorship of entertainment and cultural events drawing large minority audiences, and extensive use of billboards in minority neighborhoods.[7]

Although the cigarette industry views this as an economic issue, for the Hispanic community it is more importantly an issue of maintaining a healthy population. A number of studies have documented the diseases associated with smoking, particularly for lung cancer and cardiovascular diseases. However, several studies have indicated a low incidence of lung cancer in Hispanics, which has been attributed to the low smoking rates of Hispanics in the study samples.[9–12] If the smoking rates of Hispanics are increasing, it would be tragic if Hispanic lung cancer rates reached parity with the majority population. Evidence for a significant rise in both lung cancer and COPD (chronic obstructive pulmonary disease) mortality rates in Hispanic males has been reported recently in New Mexico for the period 1958 to 1982.[13] These data serve as an early warning signal for the future health of Hispanics, particularly if cigarette smoking rises in this population.

Given the young median age of the Hispanic population (25.1 years vs 31.9 years for the total population) and rapid growth of this population (14.6 million in 1980 and 18.8 million in 1987),[14,15] the need to prevent smoking becomes even more critical in ensuring that high rates of disease are not in its future. Trevino and Moss[16] indicate the vulnerability of the Hispanic population to high morbidity and mortality rates due to the low utilization of health services and the fact that about 30 percent of Mexican Americans (the largest Hispanic group) are without health insurance. Other health care barriers such as language, transportation, and lower education levels add to this population's vulnerability and thus increase the need for developing sound risk factor reduction programs for the prevention and cessation of smoking.

Methods

HHANES Survey Design

Details of the HHANES sample design, selection process, operational plan, and quality control procedures have been documented elsewhere.[6] As noted in the foreword section of this supplement, the HHANES was a survey of three major Hispanic subgroups in selected areas of the United States[17] conducted over a two-year period (1982–84) in three regions of the United States: Mexican Americans in

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NOTE: Author affiliations are listed elsewhere under CONTRIBUTORS.

*Haffner S, Rosenthal M, Hazuda H, Stern M: Cigarette smoking (SMK) in Mexican Americans (MA) and Anglo Americans (AA). The San Antonio Heart Study. Paper presented at the 111th annual meeting of the American Public Health Association, Dallas, TX, November 13–17, 1983.

five Southwestern states, Cuban Americans in Dade County (Miami), Florida and Puerto Ricans in the New York metropolitan area including parts of New York, New Jersey, and Connecticut. The sample design of HHANES was a stratified, multistage, probability sample of persons six months through 74 years of age. Data from all three portions of the survey are presented in this report.

The HHANES included both an interview and a three-hour medical examination conducted in a specially designed mobile examination center. Smoking history and other medical history data were collected during the household interview. The interviews were conducted in either English or Spanish, depending upon the language preference of the respondent. The vast majority of sample persons who stated they wanted to be interviewed in one language participated throughout the survey in that language. All persons ages 18 years and older responded to the questionnaires themselves unless they were physically or mentally unable to be interviewed.

The study population for this report included adults 20 through 74 years of age who met the eligibility criteria for Mexican American, Cuban American, or Puerto Rican, respectively. The analysis included those persons who completed the household interview and the medical examination. Persons not medically examined were excluded from this analysis because their smoking history responses could not be validated in the plasma cotinine study, to be described subsequently. Thus, we included only those persons for whom we could scientifically comment on the validity of the responses.

Of the 4,735 Mexican American adults ages 20–74 years who were sampled in the survey, 3,326 (70 percent) completed the household interview and medical examination, and were included in this report. Likewise, 865 (out of the sample 1,481) adult Cuban Americans and 1,220 (out of the sample 1,764) adult Puerto Ricans in this age range completed the two parts of the examination, reflecting response rates of 58.4 percent and 69.2 percent, respectively. Because these response rates are not high, as noted in the methodological paper to this supplement,[17] some caution in interpreting the results is advisable.

NHIS Survey Design

The 1985 National Health Interview Survey (NHIS) data presented in this report were analyzed by Marcus and Crane.[18] Their report was based on data collected in a continuing nationwide survey conducted on households representative of the US civilian noninstitutionalized population (one randomly selected respondent per household). Household members were interviewed by trained personnel of the US Bureau of the Census and information was obtained about the health and other characteristics of each respondent. In the 1985 NHIS, 33,045 respondents completed the smoking supplement. Standard NHIS protocols were followed in the survey, and response rates were approximately 95 percent.

In the NHIS survey, persons 17 years of age and over could respond for themselves and any responsible adult who was married or 19 years of age or older could answer questions for any other related person in the household, which may have led to some error in reporting.

Data Definitions

The cigarette smoking questions were asked of all adults during the household interview. A smoker was defined as anyone who had ever smoked at least 100 cigarettes in his/her life time. Current smokers included those among the smokers who smoked cigarettes at the time of the survey. The amount of cigarettes smoked was asked in terms of the usual number of cigarettes smoked per day, with packs equal to 20 cigarettes. The NHIS survey used similar definitions as those noted above. A validation of self-reported smoking with plasma cotinine levels was also conducted.[19] Respondents were also asked what brand of cigarettes they usually smoke.

Among the demographic variables examined in this study, education was grouped according to the highest grade or year of regular school the sample person completed. Income was defined as the total combined family income during the previous 12 months from any source. Acculturation was defined from the eight-variable scale, as described in the methodological paper for this supplement,[17] and was examined only in the Mexican American sample since the scale was not designed to assess acculturation levels of Puerto Ricans or Cuban Americans. Acculturation scores were divided into tertiles indicating low, medium, or high acculturation levels, with low acculturation indicating Mexican orientation, medium acculturation indicating bi-cultural bi-lingual, and high acculturation indicating US orientation.

Statistical Analysis

Sampling weights, specifically exam weights, were applied when calculating prevalence estimates in order to account for individual selection probabilities and adjustments for nonresponse, noncoverage, and post-stratification (Mexican American sample only) which resulted from the complex survey design used in HHANES. Each of the analytic variables was examined first using age-specific data. Since there were differences in the age composition of the three Hispanic populations, estimates were age-adjusted by means of direct standardization, using the United States 1980 Census population as the standard. These rates were computed using the program SESUDAAN.[20]

The complex survey design used in the HHANES tends to increase the variance which would have been obtained through simple random sampling.[21] As described by Kovar,[22] the design effect is the ratio of the variance of a statistic calculated from a complex sample to the variance of the same statistic calculated from a simple random sample. For HHANES, the average design effect for current smoking rates was 1.1 for Mexican Americans, .7 for Cuban Americans, and 1.3 for Puerto Ricans. Since these effects were all close to 1.0, the simple random sample variances were used in statistical testing unless noted otherwise. Two-tailed t-tests ($p < 0.05$) were used to compare point estimates. Standard errors presented in the tables were calculated using the SESUDAAN program, which takes the complex survey design into account. All data analyses were done using programs accessible through Statistical Analysis System.[23]

Since this paper is based solely on cross-sectional data, which does not allow us to investigate etiologic relationships, only descriptive statistics are presented. Multivariate analyses were not considered appropriate since the purpose of the paper was descriptive in nature.

Limitations

The cross-sectional design prohibits one from making antecedent-consequence statements, as well as from separating out any effects of selective survival on the responses. Persons who previously died because of cigarette smoking or could not participate in the examination because of illnesses due to cigarette smoking are not represented in our findings.

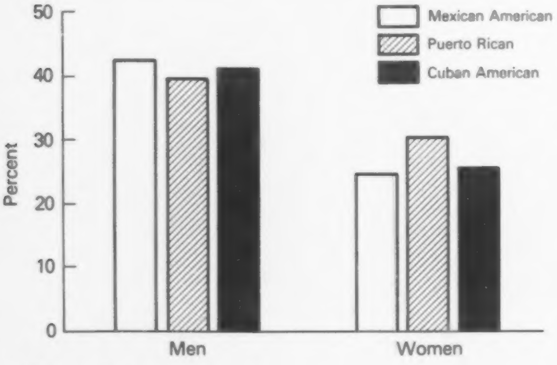
Likewise, the less than adequate response rates in the examined sample, particularly for Cuban Americans and Puerto Ricans, could have biased some of the associations. However, only slight differences were observed in the proportion of men and women who completed the examination cycle as compared to the original sample selected for study. For example, the proportion of men in the examined sample as compared to the original sample (in parentheses) among Mexican Americans 43.9 percent (47.5), Cuban Americans 43.6 percent (44.8), and Puerto Ricans 36.5 percent (39.2) did not vary by more than a few percentage points. Likewise, the percentage of older persons (ages 45–74) among the examined sample did not vary significantly from the original sample for Mexican Americans 37.7 percent (40.3), Cuban Americans 59.4 percent (58.8), or Puerto Ricans 44.2 percent (45.0).

The validation of self-reports of the number of cigarettes smoked per day with plasma cotinine levels for Mexican Americans has recently been published in this Journal.[19] The authors concluded that significant underreporting (22.6 percent) occurred in those Mexican American smokers reporting one to nine cigarettes smoked per day. There were no demographic characteristics that could be identified to characterize the underreporters. A report by Coultas *et al.*[24] among New Mexico Hispanics also found discrepancies between self-reported and validated cigarette smoking. After adjustment for cotinine and carbon monoxide levels, the age-standardized prevalence of current smokers in that study ranged from 30.9 percent (self-reported) to 39.1 percent in males and from 27.1 percent (self-reported) to 33.2 percent in females. Thus, the smoking rates reported by some Hispanics are conservative estimates of the true rates, particularly for the amount smoked.

The inconsistencies between the number of self-reported cigarettes smoked and the results of cotinine concentrations are probably due to underreporting of the quantity of cigarettes smoked per day by a proportion of Mexican American light smokers.[19] However, differences in cotinine metabolism or extremely efficient smoking are alternative explanations that cannot be ruled out by the two studies reported.[19,24]

Results

Figure 1 shows the prevalence rates of cigarette smoking among men and women in the three Hispanic groups. About 40 percent of Hispanic men were current smokers, with age-adjusted rates of 42.5, 39.8, and 41.6 percent in Mexican American, Cuban American, and Puerto Rican men, respectively. With the exception of Puerto Ricans, men were about twice as likely to smoke as women. Puerto Rican women had the highest rates of smoking, at 30.3 percent, which narrowed the sex difference in this ethnic group.



NOTE: Age-adjusted
SOURCE: National Center for Health Statistics

FIGURE 1—Age-adjusted Prevalence Rates of Current Smoking among Three Hispanic Groups in the HHANES by Sex

The age-specific smoking rates, shown in Table 1, yielded some interesting differences in the three Hispanic populations. Among men, the very high rate of smoking (50.1 percent) in young Cuban men ages 20–34 years was striking. Mexican American men had the highest rates of smoking of the three ethnic groups for ages 35–54 and 55–74 years. Among women, Puerto Ricans were smoking at higher rates than Mexican American or Cuban Americans in the two youngest ages, 20–34 and 35–54 years, with rates of 38.6 and 33.2 percent, respectively. The rates of smoking were markedly lower in the oldest age group in all three Hispanic groups, ranging from a low of 13.7 percent in Puerto Ricans to 19.5 percent in Mexican Americans.

Since comparisons of the Hispanic population with the non-Hispanic White and Black populations of the US are often made, Table 2 compares the age-specific rates, by 10-year age groups, from HHANES with smoking rates in the National Health Interview Survey conducted in 1985. The most noticeable difference was the finding that current smoking rates in the three Hispanic groups were generally higher than the rates observed for non-Hispanic Whites in NHIS, and were only slightly lower than the rates for Blacks. Exceptions were observed among Cuban Americans, where the youngest Cuban American men (ages 25–34) had markedly higher rates than Blacks (57.7 vs 45.2 percent, respectively), and where some of the older Cuban American men had about 10 percent lower rates than Blacks (ages 45–54 and 65–74).

Table 3 presents the percentage of smokers according to the amount smoked (<10, 10–19, and 20+ cigarettes per day) by age, sex, and Hispanic group. In general, the Mexican American population had a much lighter smoking pattern than the Puerto Rican and Cuban American populations. Even if we were to assume that one-fifth of the Mexican Americans who reported smoking <10 cigarettes smoked higher amounts, based on the cotinine validation study,[19] the Mexican Americans would still be lighter smokers than the other two Hispanic groups. About 52 percent of Puerto Rican male smokers and 64 percent of Cuban male smokers consumed 20+ cigarettes a day. These rates are about two times the heavy smoking rates the Mexican American rate of 34 percent rates as well as White male population cited in NHIS (37 percent).

Puerto Rican and Cuban American women smokers were also smoking large amounts of cigarettes. About one-half of Cuban women smokers and about one-third of Puerto Rican women smokers were smoking 20+ cigarettes per day. These rates are about two times those of Mexican American women in HHANES and White women in the NHIS. No marked differences from this pattern were observed for men or women when specific age groups were examined.

Since the brand of cigarette also influences health risk, we examined the usage of various cigarette brands in men and women (Figures 2 and 3). Among Hispanic men, the most popular brands used were Marlboro and Winston, together comprising 66.9, 59.0, and 46.2 percent of most often used brands among Mexican American, Puerto Rican, and Cuban American men, respectively. Contrary to expectation, ethnic brands were not used with great frequency, with the highest proportions observed for Cuban Americans (9.2 percent). Of further note was the large use of Newports by Puerto Rican men (21.0 percent) and of Benson and Hedges by Cuban American men (15.6 percent).

TABLE 1—Prevalence Rates of Current Smoking among Three Hispanic Groups in the HHANES by Sex and Age

Hispanic Ethnic Group	Sex–Age Groups							
	Men				Women			
	total (age-adjusted)	20–34	35–54	55–74	total (age-adjusted)	20–34	35–54	55–74
Mexican American								
%	42.5	40.7	44.0	43.2	23.8	22.2	29.0	19.5
± SE		1.9	2.2	3.0		1.4	1.7	2.2
Unweighted N		271	231	121		181	194	68
Puerto Rican								
%	39.8	44.7	38.8	33.6	30.3	38.6	33.2	13.7
± SE		3.9	3.7	4.7		2.9	2.6	3.0
Unweighted N		74	70	38		114	107	24
Cuban American								
%	41.6	50.1	39.2	31.6	24.4	29.2	25.0	16.2
± SE		5.3	3.8	4.3		4.3	2.9	2.9
Unweighted N		46	65	38		32	49	26

NOTE: Rates are weighted to reflect the values that would occur if the entire Hispanic population had been studied.

TABLE 2—Comparison of Current Smoking Prevalence Rates among Three Hispanic Groups in the HHANES with non-Hispanic Whites and Blacks from the 1985 Health Interview Survey

Hispanic Ethnic Group	Sex–Age Groups									
	Men					Women				
	25–34 %	35–44 %	45–54 %	55–64 %	65–74 %	25–34 %	35–44 %	45–54 %	55–64 %	65–74 %
Mexican American	42.5	43.5	44.9	44.2	41.5	23.6	30.4	26.9	21.1	16.6
Puerto Rican	41.9	42.7	33.4	32.8	35.9	41.1	34.3	30.3	12.0	16.6
Cuban American	57.7	42.3	36.8	39.0	17.7	34.7	30.1	19.3	20.6	9.7
Non-Hispanic Whites	37.1	36.3	33.4	30.1	21.2	32.6	31.4	32.9	27.4	17.3
Blacks	45.2	44.9	47.4	44.6	31.0	35.7	40.2	37.0	28.9	18.6

NOTE: Rates are weighted to reflect the values that would occur if the entire Hispanic population had been studied.

A different pattern of cigarette brand usage was seen among women in Figure 3, with Marlboro, Winston, and Salem making up a large percentage of brands used, with combined percentages of 59.7, 46.5, and 47.2 percent usage for Mexican American, Puerto Rican, and Cuban American women, respectively. Puerto Rican women favored Newports (22.0 percent) and Cuban American women favored Benson and Hedges (18.7 percent).

Socioeconomic status has often been cited as an important correlate of cigarette smoking. Table 4 and Figures 4–5 present smoking prevalence rates by income and educational level, respectively. Among men, education was inversely associated with cigarette smoking prevalence rates, for all three Hispanic groups (Figure 4). Among women, no inverse trend was observed. Women with 7–11 years of education tended to have the highest smoking rates for all three Hispanic groups. When income associations were examined in Table 4, the inverse trends were observed for Mexican American and Puerto Rican men only, with those in the lower income group smoking at greater rates than those in the higher income group. For women, similar patterns were observed and more marked for Mexican American and Cuban American groups. With the exception of the education-smoking observation in women, the data generally suggest an inverse association of socioeconomic status and smoking prevalence for both men and women.

Because income and educational levels do not reflect the degree to which the Hispanic population has adopted the US lifestyle, the association of acculturation with current smoking rates was examined for one group of Hispanics, Mexican Americans, in Figure 6. Acculturation was associated with cigarette smoking only among Mexican American women, increasing from 19 percent among Mexican-oriented women to 28 percent among US-oriented women ($p < .05$). For men, there was no clear trend for smoking and acculturation.

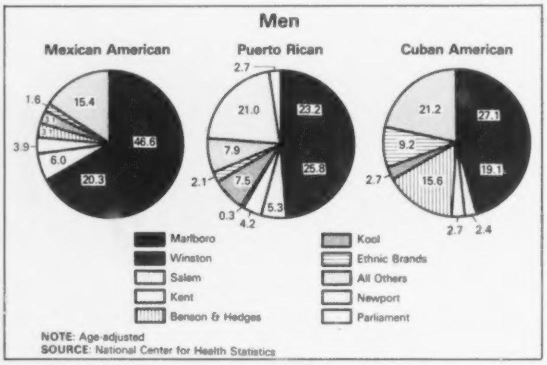
Discussion

The adverse health consequences of smoking are now widely known and accepted.[25] The health consequences of smoking that probably have the most public health impact are coronary heart disease, lung cancer, and chronic obstructive pulmonary disease. Excess rates of these diseases have not yet been observed among Hispanics as compared with the White population. If it is true, as the data show, that Hispanic smoking rates are substantially higher than those for Whites, it is likely that future Hispanic rates of these smoking-related diseases will increase. Likewise, if the estimates we have presented are conservative because of self-reporting biases, the findings of high smoking rates in Hispanics are even more dramatic than we have presented.

TABLE 3—Amount Smoked (%) by Hispanic Ethnic Group from the HHANES by Sex and Age

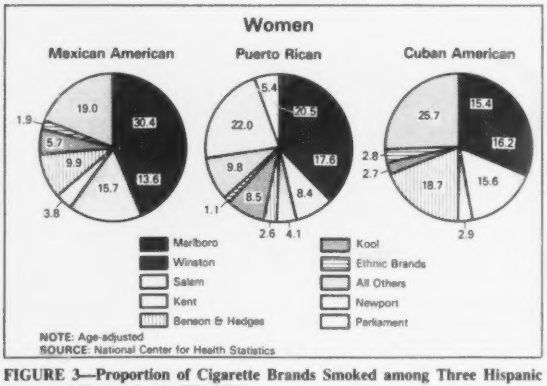
Ethnic Group/ Cigarettes Smoked per Day	Men: Total (age-adjusted) %	Men 20–34 %	Men 20–34 (±SE)	Men 35–54 %	Men 35–54 (±SE)	Men 55–74 %	Men 55–74 (±SE)	Women: Total (age-adjusted) %	Women 20–34 %	Women 20–34 (±SE)	Women 35–54 %	Women 35–54 (±SE)	Women 55–74 %	Women 55–74 (±SE)
Mexican American														
< 10	41.1	47.7	(3.0)	40.8	(3.3)	31.2	(4.2)	52.2	60.0	(3.7)	46.6	(3.6)	47.3	(6.1)
10–19	25.1	26.1	(2.6)	22.8	(2.7)	26.4	(3.9)	29.0	24.9	(3.2)	29.3	(3.2)	34.9	(5.7)
20 +	33.8	26.2	(2.7)	36.4	(3.2)	42.5	(4.5)	18.8	15.1	(2.7)	24.1	(3.1)	17.8	(4.8)
Puerto Rican														
< 10	20.7	20.1	(4.8)	22.3	(5.3)	19.5	(6.7)	32.5	28.0	(4.3)	27.8	(4.5)	45.6	(10.4)
10–19	27.0	25.5	(5.0)	25.2	(4.8)	31.8	(6.4)	32.5	39.1	(4.4)	30.9	(4.4)	24.0	(8.7)
20 +	52.3	54.4	(5.8)	52.6	(6.0)	48.7	(8.2)	35.1	32.9	(4.5)	40.7	(4.8)	30.4	(9.5)
Cuban American														
< 10	17.1	17.0	(5.4)	20.3	(5.0)	13.2	(5.6)	33.4	30.2	(8.3)	33.4	(6.9)	38.6	(9.7)
10–19	18.8	25.3	(6.3)	15.6	(4.2)	13.0	(5.3)	17.9	20.5	(6.5)	12.8	(5.0)	20.7	(7.6)
20 +	64.1	57.8	(7.3)	64.1	(6.1)	73.8	(7.2)	48.6	49.4	(9.0)	53.9	(7.6)	40.7	(10.0)

NOTE: Percentages are weighted to reflect the values that would occur if the entire Hispanic population had been studied.



NOTE: Age-adjusted
SOURCE: National Center for Health Statistics

FIGURE 2—Proportion of Cigarette Brands Smoked among Three Hispanic Groups of Men in the HHANES



NOTE: Age-adjusted
SOURCE: National Center for Health Statistics

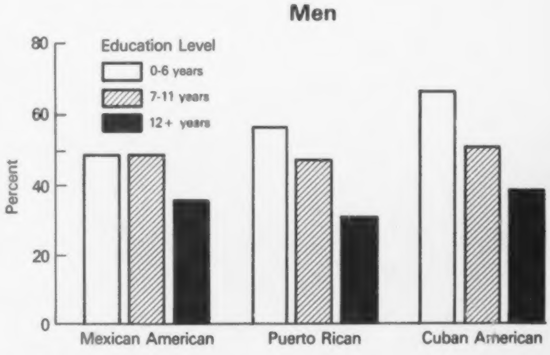
FIGURE 3—Proportion of Cigarette Brands Smoked among Three Hispanic Groups of Women in the HHANES

The HHANES results show a different picture from that reported previously.[3,6,8,26] Hispanic men in all three ethnic groups had significantly higher smoking rates than those of non-Hispanic White men in the US. In addition, Puerto Rican women had significantly higher rates of smoking than other Hispanic women, and somewhat higher rates than those of non-Hispanic White women in younger age groups. The 42.5 percent rates of cigarette smoking in the Mexican American males is quite similar to the estimates reported for men in the

TABLE 4—Age-Standardized Prevalence Rates of Current Smoking among Three Hispanic Groups in the HHANES by Sex and Income

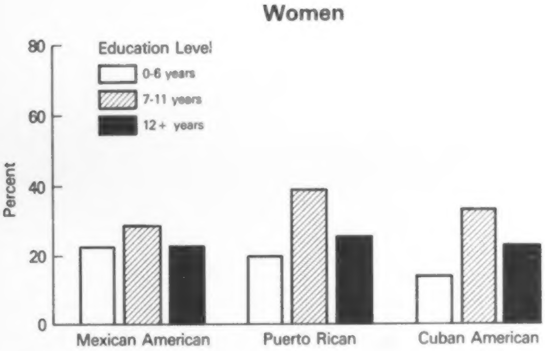
Sex and Income Group	Mexican American %	Mexican American ±SE	Puerto Rican %	Puerto Rican ±SE	Cuban American %	Cuban American ±SE
Men						
$ < 10,000	55.1	1.6	53.0	4.6	43.2	5.1
10,000 +	39.6	2.4	33.0	3.3	42.0	3.0
Women						
$ < 10,000	29.9	1.8	31.9	3.4	27.9	4.5
10,000 +	21.2	1.2	28.0	2.2	23.5	1.3

NOTE: Rates are weighted to reflect the values that would occur if the entire population had been studied. Standard errors were calculated using the SESUDAAN program (211)



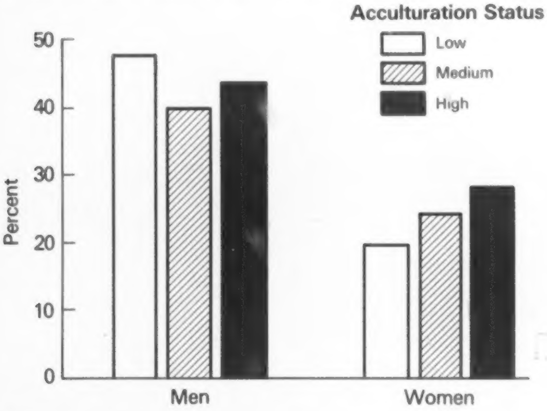
NOTE: Age-adjusted
SOURCE: National Center for Health Statistics

FIGURE 4—Age-adjusted Prevalence Rates of Current Smoking among Three Hispanic Groups of Men in the HHANES by Educational Level



NOTE: Age-adjusted
SOURCE: National Center for Health Statistics

FIGURE 5—Age-adjusted Prevalence Rates of Current Smoking among Three Hispanic Groups of Women in the HHANES by Educational Level



NOTE: Age-adjusted
SOURCE: National Center for Health Statistics

FIGURE 6—Age-adjusted Prevalence Rates of Current Smoking among Mexican American Men and Women by Acculturation Status

late 1970s in California, New Mexico, and Texas.[6,12,26–29][**] Likewise, a recent study by Markides, *et al*,[10] among Mexican Americans in San Antonio showed a 45.8 percent rate of smoking among adults ages 18 and above. Thus, it appears that the overall smoking rates in Mexican American males are not on the decline, in contrast to the trend that has been observed for other White males in this country. Estimates of smoking among Puerto Rican and Cuban American males from the 1979 and 1980 NHIS surveys were also consistent with the 40 percent rate.[5]

Comparisons of Hispanic Americans with Hispanics in the countries of origin are useful in assessing the effect of migration and acculturation on the smoking habits in the US. Data on the smoking rates in Mexico in the 1970s and 1980s have been reviewed by Silva,[30] and show a consistent range of current smoking rates in men between 40 and 45 percent.[31,†] In the same review cigarette smoking rates in women ranged between 15 and 31 percent. Thus, the overall rates for Mexican Americans in the HHANES study are comparable to rates reported in Mexico.

The best survey of Puerto Rican adults came from the Puerto Rican Heart Study, the baseline of which was conducted between 1965–68.[32] In that study, 44 percent of men ages 45–64 years were current smokers, although urban men had significantly lower rates than rural men. Since the HHANES study of Puerto Ricans was conducted in urban areas in the US, the lower rates among the urban men in Puerto Rico are comparable to the smoking rates in the HHANES Puerto Rican men. No data on Puerto Rican women are available for comparison.

Smoking data from Cuba are scarce. Only one report could be found, by Joly and Acosta,[33] who studied Cuban students in grades 10 to 12. Between 40 and 50 percent of students ages 18 and above smoked at least one cigarette per day. The authors speculated that the ultimate prevalence of smoking among Cuban adults is likely to be considerably higher than that found in students 18 years old and over. Again, these reports of high smoking rates are consistent with the 50 percent rates observed in the HHANES Cuban American males ages 20–34 years.

Although several studies have indicated that Hispanics are lighter smokers than other Whites and Blacks,[5,18] our results show that this is the case only for the Mexican American population. We have also seen from the cotinine validation study that these rates are not valid, and about one in five reported light smokers are really heavy smokers.[19] For Puerto Rican males in the New York City area and Cuban American males in the Dade County area, our data show a quite different experience, i.e., the prevalence of heavy smokers (20+ cigarettes per day) is about twice as high as the percentage of light smokers. Our findings are consistent with those of other studies in noting that the brands of cigarettes smoked by these Hispanic groups are high in tar and nicotine.[8,34][††]

The data did not allow us to examine other high-risk smoking behaviors such as inhalation, smoking the cigarette down to the tip, and average time between awakening and smoking first cigarette. However, a recent study by Marin, *et al*,[35] indicated that 49 percent of Mexican Americans (males and females) in their San Francisco sample of current smokers inhaled into the chest; 32 percent burned all the cigarette; and the average wait time to the first cigarette was 257 minutes. Men had slightly worse smoking behavior on these three factors, with 54 percent inhaling, 39 percent burning all the cigarette, and waiting 243 minutes to first cigarette. These findings suggest that there is a serious health risk among smoking Mexican Americans, as well as Cuban Americans and Puerto Ricans.

One factor that may be responsible for the high rates of smoking in the Hispanic populations is the impact of advertising on these populations. It is well recognized that cigarette manufacturers are now targeting the Hispanic and other minority populations with increased expenditures to distribute their message.[7,8,36] Direct advertising to Hispanics is not the only strategy followed by the cigarette industry. Public marketing techniques such as sponsoring meetings, publishing Hispanic directories, and offering monetary contributions to national and community Hispanic organizations to conduct highly visible programs are also used.[7] Given the young median age of the Hispanic population and its potential for tremendous growth over the next several years, it is not likely that the cigarette industry will reduce its efforts in targeting this population. Thus, the public health and medical communities as well as the Hispanic advocacy organizations must recognize this danger and respond to it appropriately.

[**] Stern MP, Haffner SM, Hazuda HP, Rosenthal M: Cardiovascular disease in Mexican Americans. Conference on Chronic disease among Mexican Americans, in Woodland, MO, December 2–6, 1984.

[†] Puenite Silva FG: Resultados de encuestas sobre el habito de fumar en tres muestras (poblacion suburbana/rural personal medico desiete centros hospitalarioso y personal de Petroleos Mexicanos). Implicaciones y consideraciones. Trabajo presentado en: II Reunion de Investigacion del Instituto Mexicano de Psiquiarria. Mexica, Noviembre de 1984.

[††] Carrillo TE, Torres IM: Risk factor prevalence among Hispanics: Socioeconomic status, diet and obesity, cigarette smoking, hypertension and alcohol. Planning for the year 2000. Sixth Biennial National Hispanic Conference on Health and Human Services. New York, September 1986.

To deal with the threat of increased smoking behavior in the Hispanic population, it will be also necessary to direct preventive efforts to specific Hispanic sub-populations. Using birth cohort data from HHANES, Escobedo and Remington[37] showed that recent cohorts of Hispanic men have reduced cigarette smoking, while recent cohorts of Hispanic women have made little progress or have actually increased their cigarette smoking. The authors recommended that intervention efforts should be especially targeted toward Cuban American and Puerto Rican women, since the proportion of smokers among successive cohorts of these women appears to have increased markedly. The association of more acculturation with cigarette smoking in women has also been reported in a recent survey of San Francisco Hispanics.[38] The authors note that this may reflect current social norms for smoking in the United States, and should be taken into consideration in the design of prevention and cessation programs for this group. In this era of health promotion and disease prevention, high rates of smoking-related diseases can and should be avoided in the Hispanic population if the proper attention and resources are devoted to prevention.

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IX. Acculturation and Marijuana and Cocaine Use: Findings from HHANES 1982–84

HORTENSIA AMARO, PHD, RUPERT WHITAKER, GERALD COFFMAN, MS, AND TIMOTHY HEEREN, PHD

Abstract: We examined the relation between acculturation and illicit drug use among Hispanics in the United States employing data from the 1982–84 Hispanic Health and Nutrition Evaluation Survey (HHANES). Across all Hispanic groups, acculturation into US society, as reflected in English language use, was associated with higher rates of illicit drug use even after sociodemographic variables such as gender, age, income, and education were considered. Significant interactions between language and education indicated that the predominant use of English was more strongly associated with marijuana and cocaine use among Mexican Americans and Puerto Ricans of lower educational attainment than among those of higher educational attainment. Significant interactions between language use and other factors such as sex, marital status, and place of birth were also associated with marijuana and cocaine use. These results suggest that the experience of acculturation, especially as it relates to drug use, is closely tied to the social and economic context in which an individual lives. [*Am J Public Health* 1990; 80(Suppl): 54–60.]

Introduction

Drug use has long been recognized as a major public health problem in the United States,[1] and increasing indicators of serious use in the last two decades have prompted a surge in research to identify patterns and risk factors for drug use in general populations.[2,3] However, these studies have generally not provided reliable epidemiological data on drug use among minorities, especially Hispanics. When data on Hispanics have been reported, most studies have failed to distinguish between Mexican Americans, Puerto Ricans, and Cuban Americans. Failure to specify results by Hispanic subgroup greatly limits the usefulness of such reports for public health efforts since, by the year 2000, Hispanics will be the largest minority group in the United States.[4] While anecdotal, clinical, and some community-based studies suggest that drug use may be a significant problem in Hispanic communities,[5–11] information on general drug use prevalence and the associated risk factors in specific Hispanic groups is essential for prevention and intervention efforts in this population.

One critical experience shared by all Hispanic subgroups is the process of acculturation or culture change resulting from continuous, first hand contact with the United States culture. Since acculturation may involve changes in attitudes, norms, and practices regarding use of illicit drugs, it is important to understand the relation of acculturation to drug use. Psychoactive drug use among Hispanics may be generally associated with increased levels of acculturation into US norms and lifestyle.[12–15] Some research has suggested[16,17] that acculturation may be more strongly associated with the use of some psychoactive substances, such as alcohol, among women than men. Whether acculturation is also more strongly associated with illicit drug use among women than men has not yet been investigated. While these findings on acculturation are suggestive, their generalizability has been limited due to the select and small samples in the reported studies. Most such studies have been limited to one specific Hispanic group or have not identified the group under study. It is also unclear from previous research whether acculturation is associated with increased drug use in all Hispanic groups. Furthermore, because acculturation is associated with improved socioeconomic status,[18] it is unclear to what extent increased drug use is prompted by economic rather than acculturative factors. Moreover, the possible interactions between acculturation and sociodemographic factors, such as sex and other demographic variables, have not yet been systematically investigated.

This paper investigates the extent to which illicit drug use among Mexican American, Puerto Rican, and Cuban American adults is associated with place of birth and language use as indicators of acculturation, and whether this relation differs across gender, age, and socioeconomic status.

Methods

The methods and sampling design used in the Hispanic Health and Nutrition Evaluation Survey (HHANES) are described in detail elsewhere.[19] Briefly, the HHANES is a complex, multistage, stratified, clustered sample of households drawn from three geographic regions (i.e., Southwestern US, Northeastern US, and Dade County, Florida) with the largest national concentrations of Hispanics. Households were screened for eligible respondents, who were asked general information about individuals in the household. The second part of the study involved physical examination and interview at a nearby mobile examination unit. Questions on drug use were included in the Adult Sample Questionnaire Supplement and administered in the interview portion of the study to individuals 12 to 74 years of age.[19]

The present analyses include respondents 20 years of age and older who self-identified as Mexican American in the Southwestern US sample, Puerto Rican in the Northeastern US sample, or Cuban American in the Dade County, Florida sample and who completed the physical examination and interview portions of the study. All respondents were asked questions on marijuana use, but only those younger than 45 years were asked questions about cocaine use. While the HHANES also included questions on use of inhalants and sedatives, data on these substances are not included in this paper due to an insufficient number of users. The participation rates for those approached, aged 20–74, were 58.4 percent among Cuban Americans, 69.2 percent among Puerto Ricans, and 70.2 percent among Mexican Americans. The non-response rate for drug questions did not exceed 1.3 percent. The Mexican American sample includes 3,303 respondents (ages 20–74) who were asked about marijuana use and 2,054 (ages 20–45) who were asked about cocaine use; the Puerto Rican sample, 1,209 and 673 respondents respectively, and the Cuban American sample, 858 and 348 respondents respectively.

NOTE: Author affiliations are listed elsewhere under CONTRIBUTORS.

Questions on drug use were limited to whether the respondent had ever used the drug in question (lifetime use) and the most recent periods of use. Due to the small number of respondents who reported use in the last month and use in the last six months, analyses were conducted only on lifetime and previous year's use (no/yes) of marijuana and cocaine. Information was also obtained on sociodemographic characteristics such as sex, age, marital status, years of education, current employment status, and yearly household income. The HHANES acculturation measure, originally developed for the Mexican American sample, is based on eight items measuring language use, ethnic identification, and birthplace of respondent and respondent's parents. An analysis of the distribution of these items for Puerto Ricans and Cuban Americans revealed that responses to some of these items are highly skewed, rendering the measure inappropriate for general use across all Hispanic groups. To develop an acculturation measure that was equally applicable to all groups, the highly skewed items were deleted and acculturation was measured by: 1) place of birth (coded as within or outside the continental US); and 2) a language index derived from the simple mean of self-ratings on language spoken, language preferred, language written, and language read (each scored English only = 1, mostly English = 2, both English and Spanish = 3, mostly Spanish = 4, Spanish only = 5), and language of interview (coded as English = 1, Spanish = 5).

Statistical Analyses

Age adjusted rates for marijuana and cocaine use: Marijuana and cocaine use rates were calculated for each Hispanic subgroup by gender. Marijuana and cocaine use rates were calculated adjusting for age by a direct standardization method. That is, each of the three Hispanic samples was weighted to reflect the age distribution of the combined weighted sample which reflects the Hispanic population in the United States. Standard errors were calculated accounting for design effects using the RTIFREQS procedure in SAS.[20] These standard errors were used to calculate 95 percent confidence intervals for drug use rates from each of the three Hispanic populations.

Marijuana and cocaine use by place of birth and language use: Marijuana and cocaine use rates by place of birth and language use were calculated. The RTIFREQS procedure in SAS[20] was employed to calculate use rate estimates and their standard errors while accounting for design effects.

Multivariate analysis to assess relationship of acculturation to drug use controlling for socioeconomic factors: To investigate whether indicators of acculturation were independently associated with drug use after the effects of socioeconomic factors were removed, a series of logistic regression analyses were conducted using LOGIST procedure in SAS.[21] The analyses were conducted separately for Mexican Americans and Puerto Ricans; the number of users in the Cuban American sample was too small for multivariate analysis to be meaningful. Use in the year prior to interview (no/yes) was chosen as the dependent variable for the analyses because it was the most recent measure of drug use with a sufficiently large sample to allow multivariate analyses. The sociodemographic variables in the model were: sex (male/female), respondent's age (continuous), yearly family income (<\$20,000/≥\$20,000), currently employed (no/yes), years of education (0–17), and marital status (measured by two dummy variables: separated/divorced/widowed [no/yes] and never married [no/yes], with married as the comparison group). The indicators of acculturation included in the model were: birthplace (continental US/not continental US), and the language index score described previously (range from 1 = Only English to 5 = Only Spanish). Proc LOGIST[1] accounts for sample weights but not design effects;[22] therefore standard errors were corrected for a design effect of 1.5 when calculating p-values and confidence limits.

Multivariate analysis to test interactions between acculturation measures and sociodemographic variables: For each Hispanic group and drug, the model described above served as the basic model to which interaction terms were added and tested one at a time. The interaction terms tested were: language use by birthplace, language use by gender, language use by age, language use by income, language use by education, and language use by marital status. Interactions between language use and birthplace and language use and each of the sociodemographic variables in the model were tested individually.

Test of the final model with all interaction terms: A final model for each drug and Hispanic group—which included all sociodemographic variables, language use, place of birth and all significant interactions as determined in the previous step—was then tested. Language use as a measure of acculturation rather than birthplace was chosen for the interactions because preliminary analyses showed this variable to be most closely associated with drug use.

Results

Sample Characteristics

Table 1 presents selected weighted sample characteristics for each Hispanic group. With the exception of the Puerto Rican sample, which was predominantly female, gender representation was approximately equal. Since drug use patterns have been shown to vary substantially by gender, the overrepresentation of females in the Puerto Rican sample of the HHANES needs to be considered in the interpretation of the findings. The Mexican American and Puerto Rican samples were younger, had a lower yearly family income, and have less formal education than the Cuban American sample. The Puerto Rican sample had a greater proportion of individuals who were unemployed and not currently married compared to the other Hispanic samples. Nearly 60 percent of Mexican Americans were born in the United States mainland, while this was true for only 21 percent of Puerto Ricans and 2 percent of Cuban Americans. Scores on the language use index indicate that Cuban Americans used Spanish more than English, Puerto Ricans were equally divided in language use, and Mexican Americans used English more than Spanish.

Drug Use

Age-adjusted rates for marijuana use: The age-adjusted rates shown in Table 2 indicate that nearly half of Mexican American and Puerto Rican men reported having used marijuana at some point in their lives, while 30.7 percent of Cuban American men reported ever using marijuana. Use of marijuana in the previous year among Hispanic men was reported by 26.6 percent of Puerto Ricans, 22 percent of Mexican Americans, and 12.5 percent of Cuban Americans.

Lifetime use of marijuana was reported by 30.8 percent of Puerto Rican women, but only 19.5 percent of Mexican American women and 11.5 percent of Cuban American women. Compared to lifetime use, use of marijuana in the previous year was reported by fewer women than men in all Hispanic groups. Compared to Mexican American (6.6 per-

TABLE 1—Selected Weighted Sample Characteristics of Hispanic Subgroup

Characteristics	Mexican American (N = 3,303)	Puerto Rican (N = 1,209)	Cuban American (N = 858)
	%	%	%
Sex			
Female	50.3	62.7	44.5
Male	49.7	37.3	55.5
Age (years)			
20–29	36.9	32.2	18.8
30–39	27.8	27.4	21.8
40–44[a]	7.5	11.4	10.7
45–49	6.9	8.8	10.9
50+	20.9	20.3	37.8
Yearly family income			
<$20,000	64.6	70.0	57.7
≥$20,000	35.4	30.0	42.3
Employed	62.2	47.6	67.5
Education (years)			
<8	33.7	21.3	27.2
8–11	24.4	31.5	19.0
12	23.8	29.5	23.7
>12	18.1	17.8	30.2
Marital status			
Never married	14.5	23.4	12.6
Married	72.0	51.6	69.3
Sep/Div/Wid	13.5	25.0	18.1
US-born[b]	59.3	21.5	2.3
Language use			
English only	25.2	18.2	4.0
Mostly English	30.3	24.0	7.9
Both equally	12.0	17.4	19.9
Mostly Spanish	19.7	31.2	42.7
Spanish only	12.9	9.2	25.6

a) The age category of 40–49 years is split into two groups because the questions on cocaine use were asked only of respondents ages 20–44.

b) Refers only to individuals born in the continental US.

TABLE 2—Age-adjusted Rates of Drug Use among Hispanic Men and Women (95 percent confidence intervals)

	Mexican American	Puerto Rican	Cuban American
Marijuana	%	%	%
Males	(N = 1,450)	(N = 440)	(N = 373)
Previous year	22.0 (19.2, 24.8)	26.6 (21.9, 31.3)	12.5 (5.9, 19.1)
Lifetime	48.8 (45.6, 52.0)	49.5 (41.3, 57.7)	30.7 (24.5, 37.0)
Females	(N = 1,853)	(N = 769)	(N = 485)
Previous year	6.6 (4.7, 8.5)	12.5 (9.8, 15.2)	3.8 (1.2, 6.4)
Lifetime	19.4 (15.8, 23.0)	30.8 (26.9, 34.7)	11.5 (5.5, 17.5)
Cocaine			
Males	(N = 903)	(N = 231)	(N = 141)
Previous year	10.8 (7.6, 14.0)	26.9 (18.7, 35.1)	17.1 (7.5, 26.7)
Lifetime	19.6 (14.6, 24.6)	40.6 (32.0, 49.2)	20.3 (8.8, 31.8)
Females	(N = 1,151)	(N = 442)	(N = 207)
Previous year	2.1 (0.8, 3.5)	13.6 (9.8, 17.4)	2.7 (0.0, 5.4)
Lifetime	6.4 (4.9, 7.9)	21.2 (16.5, 25.9)	7.2 (4.3, 11.0)

cent) and Cuban American women (3.8 percent). Puerto Rican women reported the highest rate of marijuana use (12.5 percent) in the previous year.

Age-adjusted rates for cocaine use: The age-adjusted rates for lifetime cocaine use indicate that 40.6 percent of all Puerto Rican men reported ever using cocaine and 26.9 percent reported using cocaine in the previous year. Among Mexican American men, 19.6 percent reported ever using cocaine and 10.8 percent reported using cocaine in the previous year. One-fifth (20.3 percent) of Cuban American men reported ever using cocaine and 17.1 percent said they used cocaine in the previous year.

Among women, 21.2 percent of all Puerto Rican women reported lifetime use of cocaine, while 6.4 percent of Mexican American and 7.2 percent of Cuban American women reported ever using cocaine. Use of cocaine in the previous year was reported by 13.6 percent of Puerto Rican women but only 2.1 percent of Mexican American and 2.7 percent of Cuban American women.

Marijuana and cocaine use by place of birth and language use: Figures 1 through 4 present the estimates of use and 95 percent confidence intervals (CI) for marijuana and cocaine use in the previous year by birthplace and language index. Use of marijuana and cocaine was reported more often by US-born Hispanic men and women of all groups than by those who were born outside the continental US (see Figures 1 and 2). Overall, the use of marijuana and cocaine use was highest among English-speaking Hispanic men and women than among those who were bilingual, who in turn were generally more likely to report drug use than those who were primarily Spanish-speakers (see Figures 3 and 4). As found in earlier analyses, in comparison to the other Hispanic groups, Puerto Rican men and women reported the highest use of marijuana and cocaine across language categories and birthplace. It should be noted, however, that sample sizes become very small, especially where drug use among Cuban Americans was reported by birthplace and language use, making these estimates highly unstable.

Multivariate Analysis

Table 3 presents odds ratios of socioeconomic and acculturation variables for use of marijuana and cocaine in the previous year among Mexican Americans and Puerto Ricans. The Cuban American sample of users was too small to conduct multivariate analysis. The odds ratios were

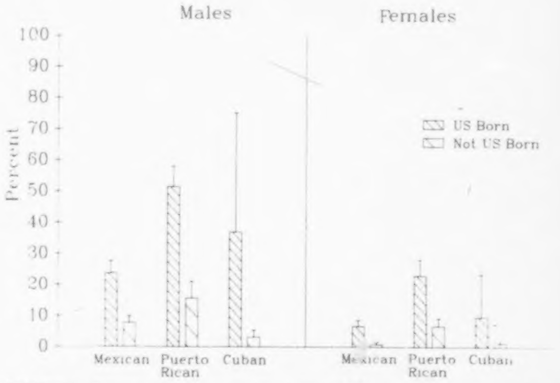

FIGURE 1—Percent of Mexican American, Puerto Rican, and Cuban American males and females who reported using marijuana in the previous year by birthplace

NOTE: Only upper 95 percent confidence limits are presented. The upper and lower limits are equidistant.

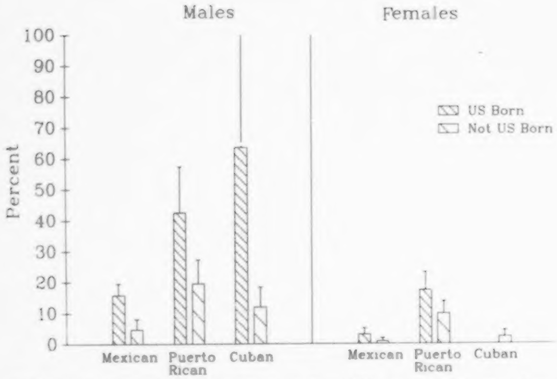

FIGURE 2—Percent of Mexican American, Puerto Rican, and Cuban American males and females who reported using cocaine in the previous year by birthplace
NOTE: Only upper 95 percent confidence limits are presented. The upper and lower limits are equidistant.

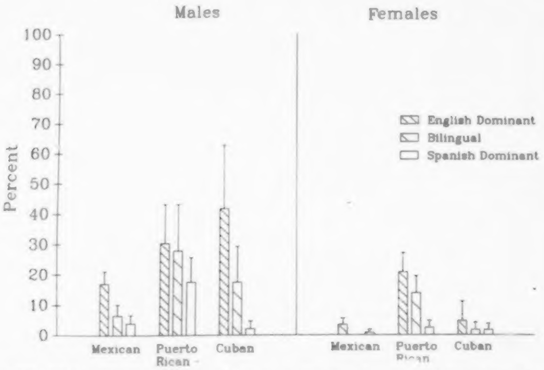

FIGURE 4—Percent of Mexican American, Puerto Rican, and Cuban American males and females who reported using cocaine in the previous year by language use.
NOTE: Only upper 95 percent confidence limits are presented. The upper and lower limits are equidistant.

developed with separate multivariate logistic regressions for each ethnic group/drug combination.

Marijuana Use

Marijuana use and acculturation effects: In both Mexican Americans and Puerto Ricans, language use was significantly associated with marijuana use in the previous year, even after sociodemographic factors were considered. The odds of using marijuana were eight times greater for Mexican Americans and five times greater among Puerto Ricans who were English-speaking than among Spanish-speakers.

Marijuana use and interaction effects: A final model, which included the significant interactions (see Methods) in addition to the basic model presented in Table 3, was tested separately for Mexican Americans ($\chi^2 = 611.84$, df = 12, $p < .0001$) and Puerto Ricans ($\chi^2 = 236.78$, df = 10, $p < .0001$).

Among Mexican Americans, marijuana use in the previous year was often reported by those who were younger ($p < .0001$), not employed ($p < .05$), separated/divorced/widowed ($p < .0002$), and never married ($p < .03$). In addition, significant interactions between language and sex ($p < .02$), language and education ($p < .05$), and language and never-married status ($p < .03$) were found. These interactions suggest that among Mexican Americans, speaking English was *more* strongly associated with a higher rate of marijuana use in the previous year among: women in comparison to men; those who have less formal education in comparison to those with more formal education; and those who were married compared to those who were not married.

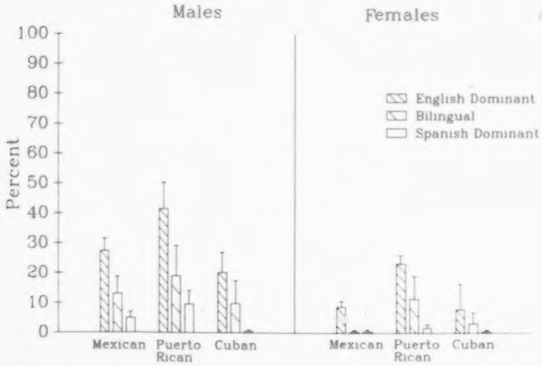

FIGURE 3—Percent of Mexican American, Puerto Rican, and Cuban American males and females who reported using marijuana in the previous year by language use.
NOTE: Only upper 95 percent confidence limits are presented. The upper and lower limits are equidistant.

Among Puerto Ricans, marijuana use in the previous year was most often reported by males ($p < .0001$), young adults ($p < .0001$), individuals with less education ($p < .05$), and those who were predominantly English-speaking ($p < .02$). One marginally significant interaction (language by education, $p < .08$) suggests that use of English among Puerto Ricans was *more* strongly associated with a higher rate of marijuana use in the previous year among those with less education than among those with more education.

Cocaine Use

Cocaine use and acculturation effects: Once sociodemographic variables were taken into account, language use was significantly associated with use of cocaine in the previous year among both Mexican Americans and Puerto Ricans (Table 3). The odds of using cocaine were 25 times greater among Mexican Americans who scored toward the English-dominant end of the language use index than among those who were Spanish-dominant. Among Puerto Ricans, English-speakers were two times more likely than were Spanish-speakers to report cocaine use in the previous year, but the association was weak.

Cocaine use and interaction effects: A final model, which included the basic model presented in Table 3 in addition to all significant interactions, was tested separately for Mexican Americans ($\chi^2 = 257.22$, df = 11, $p < .0001$) and Puerto Ricans ($\chi^2 = 158.29$, df = 12, $p < .0001$).

Among Mexican Americans, cocaine use in the previous year was most likely to be reported by those who were: male ($p < .0001$), young adults ($p < .0001$), less educated ($p < .002$), separated, divorced or widowed ($p < .006$), English-speakers ($p < .0001$), and born in the US ($p < .005$). Two significant interactions between language and birthplace ($p < .001$) and language and education ($p < .001$) suggest that use of English was *more* strongly associated with a higher rate of cocaine use among Mexican Americans who were US-born

TABLE 3—Odds Ratios (95 percent confidence intervals) of Socioeconomic and Acculturation Variables for Use of Marijuana and Cocaine in the Previous Year among Mexican Americans and Puerto Ricans, Multivariate Logistic Regressions

	Mexican Americans		Puerto Ricans	
	Marijuana	Cocaine	Marijuana	Cocaine
Variables	OR (95% CI)	OR (95% CI)	OR (95% CI)	OR (95% CI)
Sex				
Male vs. female	7.17 (4.78, 10.75)	7.52 (4.11, 13.77)	3.14 (1.93, 5.11)	2.66 (1.58, 4.49)
Age				
20 vs 40 years	8.33 (5.21, 13.32)	6.44 (2.74, 15.15)	6.03 (3.26, 11.15)	9.17 (3.63, 23.16)
Yearly Family Income				
<$20K vs ≥$20K	1.15 (0.81, 1.63)	0.75 (0.47, 1.20)	0.67 (0.40, 1.13)	0.65 (0.36, 1.17)
Employment				
Employed vs not	0.66 (0.45, 0.98)	0.94 (0.53, 1.67)	1.21 (0.72, 2.02)	1.98 (1.12, 3.50)
Education				
8 vs 16 years	0.79 (0.48, 1.29)	1.24 (0.60, 2.54)	0.91 (0.43, 1.91)	0.95 (0.42, 2.16)
Marital Status				
Sep/div/wid vs married	2.53 (1.54, 4.15)	2.75 (1.37, 5.51)	1.68 (0.89, 3.18)	3.70 (2.03, 6.74)
Single vs married	0.84 (0.56, 1.26)	1.14 (0.67, 1.95)	1.10 (0.62, 1.94)	2.55 (1.37, 4.75)
Birthplace				
US vs not US	1.54 (0.75, 3.17)	1.03 (0.50, 2.14)	1.44 (0.83, 2.49)	0.92 (0.37, 2.31)
Language Use				
English vs Spanish	8.62 (3.40, 21.88)	25.44 (6.56, 98.68)	5.32 (1.83, 15.46)	2.19 (0.69, 6.95)
ChiSquare	472.02	198.98	306.71	123.57
p	.0001	.0001	.0001	.0001

than among those who were Mexican-born and among those who had low educational attainment than among those with high educational attainment.

Among Puerto Ricans, cocaine use in the previous year was reported most often by those who were young adults ($p < .0001$), employed ($p < .04$), separated, widowed or divorced ($p < .0001$), never married ($p < .005$), and born in the US mainland ($p < .05$). Significant interactions between language and birthplace ($p < .04$) and language and sex ($p < .005$) suggest that greater use of English was *more* strongly associated with a higher rate of cocaine use among men than among women and among Puerto Ricans born on the island than those born on the US mainland. In addition, a marginally significant interaction between language and education ($p < .07$) suggests that the use of English was *more* strongly associated with cocaine use among Puerto Ricans of low educational attainment than those with higher educational attainment.

Discussion

The results indicate that the rate of marijuana and cocaine use among Hispanics, especially among Puerto Ricans, is higher than that for non-Hispanic Whites.[3] Because previous large-scale studies have not distinguished between the three largest Hispanic groups, the higher rates of drug use among Puerto Ricans may have been diluted by combining them with other Hispanics, who have drug use rates similar to those reported for non-Hispanic Whites.

It is possible that the higher rates of drug use among Puerto Ricans compared to other Hispanic groups reflect regional patterns of drug use. Some support for this interpretation is found in the data from the National Household Survey,[3] in which prevalence of lifetime and previous year marijuana and cocaine use was highest among those living in the Northeast region of the US. However, the lifetime and previous year rates for cocaine use among Puerto Ricans in the HHANES were higher than those reported in the general population of the US Northeast corridor. For example, data from the National Household Survey indicate that 8 percent of women, and 18 percent of men in the general Northeast population have used cocaine at some point in their lives. These rates were much lower than those estimated for Puerto Ricans from the HHANES (21.2 percent and 40.6 percent, respectively). Comparisons of estimates for use of cocaine in the previous year follow a similar pattern, suggesting that use of cocaine cannot be attributed solely to the geographic location.

The results also indicate that acculturation into US society, as reflected by language use, was accompanied by a higher prevalence of illicit drug use. This relationship holds true even when sociodemographic variables are taken into account. Yet, the relative strength of the association between acculturation measures and drug use varied across Hispanic subgroup and by drug. While the cross-sectional design of this study does not allow for an analysis of the process of acculturation, the results lead us to believe that it is important to develop research to further investigate the nature of the acculturation process among Hispanic subgroups and its relationships to drug use.

One possible factor which influences the impact of acculturation on drug use is the social context of drug use in the native country. For example, it is possible that drug availability and norms in Puerto Rico are closer to those in the US mainland than to those in Mexico or Cuba. This interpretation is supported by previous reports of drug use in Puerto Rico.[23,24] Thus, in order to better understand the factors which impact drug use among Hispanics living in the US mainland, research is needed on social norms and drug use practices in the countries from which Hispanics originate.

The results also suggest that the relation of acculturation to drug use differs across education levels. Acculturation, as measured by language use, had the strongest relation to drug use among those who were least educated. Mexican Americans and Puerto Ricans who were highly acculturated, but had not enjoyed access to educational resources of American society, were most likely to report marijuana and cocaine use. Similar findings indicating that high acculturation is associated with alcohol use among poor Hispanic men have been reported in the literature.[17,25,26] Acculturation, when accompanied by poverty and/or lack of access to education, may resemble what Berry termed "marginalization."[27] In this type of acculturation, an individual loses essential features of his/her culture and at the same time has not entered the larger society. This type of acculturation results in feelings of alienation and loss of identity, is highly stressful, and places an individual at high risk for poor mental health outcomes.[27]

Our findings also suggest that, in some cases, the relationship of acculturation to drug use varies by sex. Once socioeconomic factors were considered, the gender gap in marijuana use was relatively small among Mexican Americans who speak primarily Spanish. This contrasts with the rates of drug use among those who were predominantly English-speaking, for whom the rates were much higher among men than among women. Yet, the relative difference in drug use between Spanish- and English-speakers was larger among women than men. This finding is consistent with reports in the alcohol literature, which suggest that acculturation is more strongly associated with alcohol use among Mexican American women than among Mexican American men.[16,17] However, among Puerto Ricans, acculturation was more strongly associated with cocaine use among men than among women. The findings indicating that acculturation is differentially associated with drug use across educational levels and gender leads us to suspect that acculturation, as it relates to drug use, is integrally affected by the social and economic context in which an individual becomes socialized and lives. The impact of gender roles and educational attainment on drug use and their relation to acculturation among Hispanics deserves empirical investigation.

HHANES, although the first large scale survey to provide information on drug use among various Hispanic groups in the US, is limited in several ways. First, the measures of acculturation used in the HHANES were developed for Mexican Americans and not for Puerto Ricans and Cuban Americans. Second, measures of drug use employed in the HHANES do not allow assessment of the degree or context of drug use or the impact of drug use on functioning. Third, the HHANES employed self-reported measures of drug use, which are vulnerable to multiple threats to validity.[28] Respondents tend to underreport socially undesirable behaviors. What is deemed undesirable behavior may differ among women and men and among those who are not acculturated compared to those who are highly acculturated. If gender roles and culture determine norms regarding drug use, then the validity of self-reported drug use may also vary by these factors. Even with these limitations, the findings have important implications for the study of drug use among Hispanics because they provide evidence of a relationship between acculturation and illicit drug use. The findings also underscore the need for a better understanding of the process of acculturation among Hispanics and other immigrant groups.

The present findings also indicate the need for further studies of drug use among Hispanics, especially Puerto Ricans on the mainland and on the island. Because of the close political, economic, and legal ties between Puerto Rico and the United States mainland, and because of the continuous migration to and from Puerto Rico, drug availability and norms of drug use in the continental US may have a direct effect on drug use among Puerto Ricans living on the island. Yet, Puerto Rico is not usually included in national studies of drug use, leaving a critical gap in knowledge about a substantial Hispanic population in the US. Further, public education campaigns and prevention efforts to reduce initiation into drug use have not typically included Puerto Ricans on the island. Policy initiatives are needed to develop and implement prevention programs aimed at the Puerto Rican population on both the island and in the mainland. Such initiatives should include ongoing mechanisms for gathering information on drug use among Puerto Ricans.

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X. Generational Differences in Perinatal Health among the Mexican American Population: Findings from HHANES 1982–84

SYLVIA GUENDELMAN, DSW, JEFFREY B. GOULD, MD, MPH, MARK HUDES, PHD, AND BRENDA ESKENAZI, PHD

Abstract: Data from the Hispanic Health and Nutrition Examination Survey (HHANES) were used to examine a profile of social, medical, and behavioral characteristics associated with low birthweight (LBW) and miscarriages in first and second generation Hispanics of Mexican descent. The percentage of LBW was 5.3 and of miscarriages was 12.7. LBW rates were higher for second generation primipara and multipara compared with first generation women. Using multivariate logistic regression techniques and adjusting for complex design effects, generation was found to be a significant predictor of LBW but not of miscarriages. The findings support existing evidence that a Mexican cultural orientation protects first generation, Mexico-born women against a risk for LBW. However, the findings do not show significant effects of generation on miscarriages, suggesting that cultural effects are not consistent for all pregnancy outcomes. Furthermore, we suggest that the higher rates of LBW in second generation women are not due to a higher rate of miscarriages as has been hypothesized. [*Am J Public Health* 1990; 80(Suppl):61–65.]

Introduction

The Mexican-origin population in the United States has been growing sharply since 1970, largely as a result of high fertility rates and immigration. Despite the high birth rates of Hispanic women, little is known about their perinatal outcomes and their associated risk factors. Existing evidence, mainly derived from birth and death certificates, suggests that Mexican-origin women have birth outcomes similar to those of White non-Hispanic populations.[1–9]

Recent findings also suggest that birth outcomes are more favorable among first generation women born in Mexico than among second or successive generations of US-born women of Mexican descent. According to a 1985 nationwide study, only 5.0 percent of the infants of Mexico-born mothers were of low weight, compared to 6.3 percent of infants of US-born Mexican origin mothers and 5.7 percent of white non-Hispanic mothers.[7] Similar differential rates by birthplace have been found in California, the state with the largest concentration of Mexican descendants,[1] and most recently by Scribner and Dwyer[10] in a study of the Hispanic Health and Nutrition Examination Survey (HHANES).

The lower rates of low birthweight (LBW) of first generation mothers is intriguing and difficult to explain according to evidence derived from birth certificates.[1,4–7] Compared to US-born mothers of Mexican descent, foreign-born, first generation women have a lower socioeconomic status, a higher percentage of women over 35 years of age, and less adequate prenatal care.[2–4] On the basis of these risks, one would expect first generation mothers to have an increased risk of LBW. Recent evidence from the HHANES challenges this assumption. According to Scribner and Dwyer women with a Mexican cultural orientation have a reduced risk of LBW, even after controlling for sociodemographic characteristics and smoking.[10] A Mexican orientation serves as a marker for a lifestyle that is protective against the negative influences on pregnancy outcomes.[10,11]

A few other studies have suggested that the lower LBW rate in Mexico-born women may be related to a high rate of miscarriages which eliminate weaker fetuses.[12,13] Approximately 10 percent to 15 percent of pregnancies identified by conventional criteria end in clinically recognizable spontaneous abortions.[14] Miscarriage/stillbirth rates for Mexican American farmworkers range from 24 percent to 31 percent, depending on the target population and method of ascertainment.[15,16]

In this study, we use the HHANES data base to examine the acculturation and miscarriage hypotheses regarding more favorable LBW rates in Mexico-born compared with US-born women of Mexican origin. For this purpose we first explore a wide array of social, behavioral, and medical variables associated with LBW and miscarriages in first and second generation women of Mexican descent. We subsequently investigate the effects of generation on LBW, controlling for social, behavioral, and medical factors in separate predictive models for primipara and multipara. The use of separate models is based on findings by Eisner, *et al*,[17] indicating that multipara have a lower percent LBW than primipara and different factors predict LBW in each group. Finally, controlling for the other maternal characteristics in a multivariate regression model, we examine whether generational status also has an effect on miscarriages.

Methods

Sample

Data were obtained from the Medical History section of the HHANES which is a health survey conducted in 1982–84 of three Hispanic sub-populations (Mexican American, Puerto Rican, Cuban American)[18] living in selected areas of the United States.[18] The survey contains information on approximately 8,500 Mexican Americans residing in the southwestern US. Representatives were selected using a multistage, stratified area probability sample of households and residential clusters. All adult women of Mexican birth or origin, were initially included in this study if they were interviewed in the household and examined in a mobile examination unit and were between the ages of 16 and 55. We excluded women over age 55 due to concerns regarding recall accuracy. Out of 1,518 participants, 128 (8.4 percent) were excluded due to missing values of LBW, miscarriages, and/or generation which are the key variables of interest. Thus the final analytic sample consisted of 1,390 respondents.

Definition of Variables and Data Analysis Technique

Our two outcome variables are a history of LBW and of miscarriages (0 = no; 1 = yes) derived from the following HHANES questions:

- For *multipara*, "How many of your children (who were born alive) weighed less than 5½ pounds (2,500

Note: Author affiliations are listed elsewhere under CONTRIBUTORS.

grams) at birth?" and "What is the total number of live births you have had?"

- For *primipara*, "Did this (the only) child weigh less than 5½ pounds (2,500 grams) at birth?"
- Other questions asked of *all women* were "What is the total number of miscarriages you have had?" and "How many times have you been pregnant?"

Generation, which is determined by the birthplace of the subjects, was selected as the primary independent variable. To allow for sufficient sample size, the original three-generation coding given in the data tape (first generation born in Mexico; second generation not born in Mexico, one or both parents born in Mexico; third or higher generation, neither parent born in Mexico) was collapsed into two groups: 1) subjects born in Mexico whose parents were born outside the US, and 2) subjects born in the US who had at least one parent born in Mexico or of Mexican descent.

Other social, behavioral, and medical variables were initially screened if they have been reported in the literature as significantly related to birthweight or miscarriages in non-Hispanic populations.[19,14] Only those variables that were at least moderately associated ($p \le .20$) with LBW in pairwise analyses and/or with miscarriages in separate pairwise analyses (adjusting for number of live births and number of pregnancies, respectively) were included for further study.

The *social variables* considered in the predictive models for LBW and miscarriages are: 1) Medicaid coverage in the last 12 months prior to the interview; 2) highest grade of regular school completed; 3) family income which is the combined income obtained in the last 12 months from jobs, public assistance, rents, interests or other sources; 4) marital status which refers to whether a spouse is living in the home; and 5) age in years at the time of the interview.

The *behavioral factors* are: 1) current drinker status determined by asking of non-abstainers whether they have had a drink in the last 28 days; 2) current smoker status determined by the question "Do you currently smoke cigarettes?"; and 3) interval, in years, since the last Pap smear test, recoded into four ordinal categories ranging from last two weeks to never.

Medical factors include: 1) self-reported physician-diagnosed blood pressure or hypertension; 2) Quetelet index, i.e., a body mass index based on self-reported weight at age 25 and height (weight/height^2). On the basis of the pairwise analyses described above, three additional variables were included in the model for miscarriages, namely: self-reported physician diagnosed diabetes, heart problems, and self-perceived health status ranging from very good to poor.

Differences between the two generational groups for each independent and outcome variable were examined by performing Student's two sample t-tests for the continuous variables and chi-square tests for the categorical variables. Since the HHANES is a multistage, stratified probability sample of clusters of persons in area-based segments,[18] estimates were weighted using the "examined weights" and complex design effects were taken into account using the Super Carp program.[20] To facilitate presentation, only the weighted values and statistics are presented in the generational profile in Table 1.

Logistic regression was used (SAS Logist procedure)[21] to examine the effects of generation on each outcome variable controlling for the social, medical, and behavioral factors studied. Only cases which had complete information on all these variables were included in the analyses (n = 1,078). We ran simultaneous logistic regressions for all women and then separate models for primipara and multipara. From these results we estimated odds ratios and their 95 percent confidence intervals. For continuous or ordinal predictors, standardized odds ratios are estimated. These are presented in Tables 2–5. To check whether the statistical significance of some variables might be affected by collinearity with other variables, the simultaneous logistic procedure was followed

TABLE 1—Maternal Characteristics by Generation for Hispanic Women, HHANES 1982–84

	First Generation			Second Generation and Beyond			
Characteristics	N = 543	Weighted Mean or %*	S.E.	N = 847	Weighted Mean or %*	S.E.	p**
Social							
Education, M in years	543	M = 7.1	0.2	847	M = 10.6	0.3	.000
Income, M in $1,000 increments	543	M = 14.1	0.3	847	M = 15.8	0.4	.009
Medicaid last 12 months	541	5.9%	1.4	842	8.6%	1.1	.168
Married	541	75.8%	2.3	847	70.0%	2.0	.094
Age, M in years	543	M = 34.1	0.5	847	M = 34.3	0.4	.763
Live births, M number	543	M = 3.3	0.8	847	M = 3.1	0.4	.827
Behavioral							
Current smoker	528	21.1%	1.3	798	28.1%	1.6	.009
Current drinker	542	6.0%	1.0	847	12.4%	1.3	.005
Never had a Pap smear	538	11.1%	1.3	832	3.3%	1.1	.002
Biological							
History of high blood pressure	543	21.4%	2.2	847	22.4%	1.9	.740
Quetelet index at age 25	446			801			
Low Quetelet index		30.8%	1.6		28.6%	1.4	
Medium Quetelet index		56.8%	2.3		55.7%	1.7	.517
High Quetelet index		12.5%	2.3		15.7%	2.7	
History of diabetes	543	5.1%	0.5	847	8.0%	0.4	.002
History of heart problems	543	2.4%	0.6	846	4.8%	0.8	.043
Perceived health status, M	542	M = 3.28	0.03	847	M = 2.83	0.04	.000
Pregnancies, M number	543	M = 3.9	0.7	847	M = 3.7	0.6	.834

SOURCE: Hispanic Health and Nutrition Examination Survey, 1982–1984.

*Complex design effects also taken into account.

**p values represent the significance level for the first and second generation comparison in the weighted analysis.

up with a backwards stepwise regression.[21] Data in the models were weighted according to the scaled weighting procedure proposed by the National Center for Health Statistics, where the sum of the weights is equal to the observed sample size.[22] The results on Tables 2–5 were also examined adjusting for complex design effects using the estimated average design effect of 1.5 as described by Delgado, et al.[23]

Results

Generational Profiles of Maternal Characteristics

The two generational groups were similar with respect to age at interview (34.1 years for the first and 34.3 years for the second) and the proportion of primiparous women (19.7 percent in the first and 20.8 percent in the second). The findings in Table 1 indicate that a variety of significant changes take place between the first and second generation or beyond. Second generation women have a significantly higher educational attainment than first generation women. On the average they have 3½ more years of formal schooling. Family income is also significantly higher in the second generation despite the fact that, compared with the first generation, there are more women living without a spouse at home. From a behavioral standpoint, far more second generation women are smokers and drinkers than first generation women. However, significantly fewer second generation women have never had a Pap smear compared with those in the first. No significant differences in high blood pressure or Quetelet index were found between the two generations, but second generation women had significantly higher proportions of physician-diagnosed diabetes and heart problems. Despite reporting more medical problems, second generation women perceived their health status significantly better than first generation women.

Generational Effects on Low Birthweight

The percentage of low birthweight out of the total live births for Mexican origin mothers is 5.3 (S.E. = 0.04). This estimate is similar to the rates based on birth certificate data reported by Williams[1] (5.3 percent) and Ventura[5] (5.6 percent) and higher than those reported by Scribner and Dwyer[10] (4.8 percent) using an HHANES sample of older women up to age 74. Similar to Scribner and Dwyer, we found that LBW rates differ significantly for the two generational groups. Out of the total live births for the first generation 3.9 percent (S.E. = 0.7) are of LBW, whereas out of the total live births for the second generation, 6.1 percent (S.E. = 0.7) are of LBW. The generational effects on LBW are significant, even after adjusting, in a logistic regression model, for the social, behavioral and medical characteristics. The risk of LBW is 1.73 times higher (95% CI = 1.11, 2.71) in the second generation compared with the first (Table 2). This result remains significant even after adjusting for complex design effects.

We examined the generational differences in birthweight outcomes in the two parity groups and found that, in the first generation, the LBW rate was 6.3 percent (S.E. = 1.8) in primipara and 3.8 percent (S.E. = 0.05) in multipara. In the second generation, primipara also had higher rates: 9.7 percent (S.E. = 2.4) compared with 5.9 percent (S.E. = 0.06) in multipara. These estimates indicate that although the percentage of LBW increases in the second generation, in both generations primipara have LBW rates that are almost 1.5 times higher than multipara.

TABLE 2—Adjusted Odds Ratios[+] for Low Birthweight (n = 1078)

Variables	Adjusted Odds Ratios[*]	95% CI
Generation (0 = first, 1 = second or beyond)	1.73	1.11, 2.71
Number of live births	1.37	1.13, 1.67
Age at interview	1.13	0.91, 1.41
Marital status (0 = no spouse home; 1 = spouse at home)	0.92	0.58, 1.47
Years of education	0.93	0.75, 1.15
Income	0.93	0.75, 1.14
Medicaid coverage (0 = no; 1 = yes)	0.83	0.39, 1.78
Current drinker status (0 = abstainer; 1 = drinker)	0.76	0.41, 1.40
Current smoker status (0 = no; 1 = yes)	1.58	1.05, 2.38
Interval since last Pap smear	1.01	0.84, 1.21
Low Quetelet index at age 25 (0 = no; 1 = yes)	1.46	[**]
High Quetelet index at age 25 (0 = no; 1 = yes)	1.26	
History of high blood pressure (0 = no; 1 = yes)	1.25	0.82, 1.91

SOURCE: Hispanic Health and Nutrition Examination Survey; 1982–84

[+] Logistic regression analysis techniques were used to obtain the adjusted odds ratios.

[*] For the continuous and ordinal variables, live births, age at interview, education, income and interval since last pap smear, standardized odds ratios are presented. (The odds ratio of LBW for the increase of one standard deviation of the predictor.)

[**] To evaluate the Quetelet index (a nominal variable with 3 categories), a χ^2 test statistic with 2 D.F. was obtained from the difference of two log likelihood functions and was found to be not statistically significant.

TABLE 3—Adjusted Odds Ratios[+] for Low Birthweight among Primiparous Women (n = 187)

Variables	Adjusted Odds Ratios[*]	95% CI
Generation (0 = first, 1 = second or beyond)	4.08	0.81, 20.45
Age at interview	0.97	0.53, 1.77
Marital status (0 = no spouse home, 1 = spouse at home)	1.23	0.31, 1.51
Years of education	0.98	0.51, 1.91
Income	1.32	0.68, 2.59
Medicaid coverage (0 = no; 1 = yes)	0.30	0.02, 4.18
Current drinker status (0 = abstainer; 1 = drinker)	0.93	0.78, 1.11
Current smoker status (0 = no; 1 = yes)	0.62	0.16, 2.43
Interval since last Pap smear	1.87	1.09, 3.20
Low Quetelet index at age 25 (0 = no; 1 = yes)	1.28	[**]
High Quetelet index at age 25 (0 = no; 1 = yes)	1.06	
History of high blood pressure (0 = no; 1 = yes)	5.07	1.59, 16.20

SOURCE: Hispanic Health and Nutrition Examination Survey; 1982–84

[+] Logistic regression analysis techniques were used to obtain the adjusted odds ratios.

[*] For the continuous and ordinal variables, standardized odds ratios are presented.

[**] To evaluate the Quetelet index (a nominal variable with 3 categories), a χ^2 test statistic with 2 D.F. was obtained from the difference of two log likelihood functions and was found to be not statistically significant.

As shown in Table 3, we found that after adjusting for the potential confounders included in the logistic regression model, the risk of LBW is approximately four times higher for second than first generation primiparous women (OR 4.08; 95% CI = 0.81, 20.45). The risk factor, however, was not statistically significant at the p = .05 level, possibly due to small sample size. The risk for LBW is five times higher for primiparous women with a physician-diagnosed history of high blood pressure (OR 5.07; 95% CI = 1.59, 16.20) and it almost doubles when the interval since last Pap smear is greater than one standard deviation from the mean (OR 1.87; 95% CI = 1.09, 3.20). After adjusting for complex design effects, the only significant predictor that remained was blood pressure.

The effect of generation on LBW was significant for multiparous women. The findings in Table 4 show that the risk of LBW is almost double for second generation women (OR 1.69; 95% CI = 1.05, 2.84) and for current smokers (OR 1.79; 95% CI = 1.16, 2.76). Smoking remained a significant predictor even after adjusting for complex design effects. Results similar to those obtained with the logistic regression techniques were obtained when doing backwards stepwise regressions for primipara and multipara.

Generational Effects on Miscarriages

The percentage of miscarriages out of the total number of pregnancies for Mexican origin mothers is 12.7 percent (S.E. = 0.59). This rate falls within the expected range for spontaneous abortions in the mainstream population.[14] Furthermore, the rates for the two generational groups are not significantly different. Out of the total number of pregnancies for the first generation, 13.1 percent (S.E. = 0.99) have had miscarriages. For the second generation the percentage of miscarriages is 12. 5 percent (S.E. = 0.60).

Using logistic regression, we examined whether generation becomes a significant predictor of miscarriages after adjusting for the social, behavioral and medical factors listed in Table 5. Generation was not a significant predictor of miscarriages (OR 0.81; 95% CI = 0.57, 1.15). The adjusted odds ratios reported in Table 5 indicate that miscarriages increase 1.25 times with every increase in one standard deviation in education (95% CI = 1.04, 1.51); it increases 1.47 times (95% CI = 1.79, 1.22) with every decrease in one standard deviation in age and 1.43 times among women without a spouse living at home (95% CI = 2.13, 0.97). Education and age remained significant after adjusting for complex design effects.

Since Scribner and Dwyer[10] reported that generation was not as strong a predictor of pregnancy outcome as acculturation, we examined the role of acculturation in predicting miscarriages. Like Scribner and Dwyer, we looked at the effect of the acculturation index with and without including generation and found that it was not a significant predictor of miscarriages.

TABLE 4—Adjusted Odds Ratios[+] for Low Birthweight among Multiparous Women (n = 892)

Variables	Adjusted Odds Ratios*	95% CI
Generation (0 = first, 1 = second or beyond)	1.69	1.05, 2.84
Number of live births	1.37	1.11, 1.68
Age at interview	1.18	1.05, 1.33
Marital status (0 = no spouse home, 1 = spouse at home)	0.86	0.52, 1.44
Years of education	0.92	0.73, 1.16
Income	0.89	0.79, 1.11
Medicaid coverage (0 = no; 1 = yes)	0.97	0.43, 2.20
Current drinker status (0 = abstainer; 1 = drinker)	0.79	0.41, 1.52
Current smoker status (0 = no; 1 = yes)	1.79	1.16, 2.76
Interval since last Pap smear	0.89	0.72, 1.09
Low Quetelet index at age 25 (0 = no; 1 = yes)	1.46	**
High Quetelet index at age 25 (0 = no; 1 = yes)	1.38	
History of high blood pressure (0 = no; 1 = yes)	0.95	0.60, 1.51

SOURCE: Hispanic Health and Nutrition Examination Survey; 1982–84

[+] Logistic regression analysis techniques were used to obtain the adjusted odds ratios.

*For the continuous and ordinal variables, standardized odds ratios are presented.

**To evaluate the Quetelet index (a nominal variable with 3 categories), a χ^2 test statistic with 2 D.F. was obtained from the difference of two log likelihood functions and was found to be not statistically significant.

TABLE 5—Adjusted Odds Ratios[+] for Miscarriages (n = 1078)

Variables	Adjusted Odds Ratios*	95% CI
Generation (0 = first, 1 = second or beyond)	0.81	0.57, 1.15
Pregnancy number	3.92	3.09, 4.96
Age at interview	0.68	0.56, 0.82
Marital status (0 = no spouse home, 1 = spouse at home)	0.69	0.47, 1.03
Years of education	1.25	1.04, 1.51
Income	1.15	0.97, 1.37
Medicaid coverage (0 = no; 1 = yes)	0.96	0.52, 1.75
Current drinker status (0 = abstainer; 1 = drinker)	0.99	0.60, 1.61
Current smoker status (0 = no; 1 = yes)	1.11	0.78, 1.58
Interval since last Pap smear	0.91	0.78, 1.06
Low Quetclet index at age 25 (0 = no; 1 = yes)	1.13	**
High Quetelet index at age 25 (0 = no; 1 = yes)	0.96	
History of high blood pressure (0 = no; 1 = yes)	1.23	0.86, 1.76
History of diabetes (0 = no; 1 = yes)	1.25	0.70, 2.24
History of heart problems (0 = no; 1 = yes)	1.07	0.52, 2.20
Perceived health status	1.09	0.93, 1.29

SOURCE: Hispanic Health and Nutrition Examination Survey; 1982–84

[+] Logistic regression analysis techniques were used to obtain the adjusted odds ratios.

*For the continuous and ordinal variables, standardized odds ratios are presented.

**To evaluate the Quetelet index (a nominal variable with 3 categories), a χ^2 test statistic with 2 D.F. was obtained from the difference of two log likelihood functions and was found to be not statistically significant.

Nor was generation a significant predictor of miscarriages when logistic models were constructed separately for primipara and multipara (primipara: OR 0.76; 95% CI = 0.31, 1.87; multipara: OR 0.91; 95% CI = 0.62, 1.34). However, for primipara, an older age at interview (OR 1.74; 95% CI = 1.17, 2.57) and Medicaid coverage (OR 3.30 95% CI = .93; 11.69) was positively associated with miscarriages, whereas in multipara age and socioeconomic factors were inversely associated with miscarriages (age OR 0.64; 95% CI = 0.52, 0.78; income OR 1.17; 95% CI = 0.97, 1.42; education OR 1.22; 95% CI = 1.00, 1.49).

Discussion

The findings from this study show that for the Mexican origin population in the United States the percentage of LBW was 5.3 and the percentage of miscarriages was 12.7. Whereas the incidence of LBW in this population is lower than that for the general population, the rate of miscarriages seems similar to what is expected for the mainstream population.[5,8,14]

Mexican Americans are a heterogeneous group and our analysis demonstrates that wide variations are found according to generation and parity. US-born Mexican Americans are at 60 percent higher risk for LBW than Mexico-born women. Furthermore, the rate of LBW among US-born primipara at 9.7 percent begins to approximate that of the Black population.[19]

The findings, based on a rich array of variables, further demonstrate that across generations, significant changes take place in certain bio-social-behavioral characteristics of mothers which are associated with pregnancy outcomes (Table 1). While second generation Mexican American women show a higher educational and income status and more use of health care relative to their first generation counterparts, certain behavioral characteristics such as smoking and drinking behaviors actually deteriorate in the second generation. The second generation also has higher rates of physician-diagnosed diabetes and heart problems. However, this could be a result of improved access to care rather than to poor health status.

Our multivariate models indicate that generation is a significant predictor of LBW but not of miscarriages. This evidence points to two important findings: first, even after controlling for other factors, US-born Mexican origin women of the second generation or beyond have a higher likelihood of giving birth to LBW infants; and second, these patterns are not consistent for all pregnancy outcomes.

According to the acculturation model, the behaviors or lifestyles that protect women against adverse pregnancy outcomes are associated with a Mexican cultural orientation which predates pregnancy. If this were the case, we would have expected that generation (as a proxy for acculturation) would have been a significant predictor, not only of LBW but also of miscarriages. Evidence from this study does not support this conclusion. The findings show no significant effects of generation or the acculturation index on miscarriages. Rather than cultural factors, the results suggest that socioeconomic, genetic, and unaccounted medical factors are better predictors of miscarriages in Latinas.

Furthermore, the findings do not support the hypothesis that increased fetal losses explain the more favorable LBW outcomes in first generation Mexican-Americans.[12,13] The 13.1 percent miscarriage rate in the first generation was very similar to the 12.5 percent found in the second generation.

However, the possibility of reporting bias cannot be discounted in this study. Underreporting miscarriages among first generation women could occur as a result of less education which leads to less body awareness or less likelihood of recognition of miscarriages. Furthermore, compared with US-born, Mexico-born women are more likely to seek prenatal care late or none at all,[5] and thus are less likely to receive a medical diagnosis if miscarriages occur. Conversely, because induced abortion in Mexico is illegal, Mexico-born women could be misreporting induced abortion as miscarriages to make it more acceptable.[24] The possibility of measurement error also arises given the potential ambiguity concerning what events actually constitute a miscarriage. We suspect that miscarriages may be more prone to reporting bias than LBW, particularly among less educated women. Hence, further studies are needed to confirm the validity of the fetal loss hypothesis.

Because the HHANES is a cross-sectional, retrospective survey, variables do not necessarily measure the conditions or behaviors at the time of birth, hence possibly washing away some of the effects. In this study, the recall bias was as long as 40 years and no attempts at verifying the accuracy of the miscarriage or LBW information given to the mother at the time of delivery were made in the HHANES. Nevertheless, some of the traditional risk factors e.g. smoking for LBW and e.g. age for miscarriages are corroborated using the HHANES data. Furthermore, the results of LBW are consistent with those found by Williams, *et al*,[1] in California. The latter found that the lower rates of LBW among Mexico-born compared with US-born women of Mexican origin could not be the result of systematic underreporting by Mexican women.

In conclusion, the findings from the HHANES suggest that cultural explanations of differential pregnancy outcomes in women of Mexican origin must be restricted to LBW. The evidence strongly argues for the need to include generation as a risk factor when doing studies of LBW in Mexican-Americans. Further studies are also needed to assess the extent to which prematurity or intrauterine growth retardation determine pregnancy outcomes in each generation. This is a question that cannot be addressed in the HHANES.

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XI. The Prevalence of Total Tooth Loss, Dental Caries, and Periodontal Disease among Mexican Americans, Cuban Americans, and Puerto Ricans: Findings from HHANES 1982–1984

AMID I. ISMAIL, BDS, MPH, DRPH, AND SUSAN M. SZPUNAR, RDH, DRPH

Abstract: This paper describes the prevalence of total tooth loss, dental caries, and periodontal disease in 2,226 Puerto Ricans, 1,192 Cuban Americans, and 5,983 Mexican Americans, ages five to 74 years, who were examined during the 1982–84 Hispanic Health and Nutrition Examination Survey (HHANES). The prevalence of total tooth loss was 2.60, 6.10, and 2.80 percent among Mexican Americans, Cuban Americans, and Puerto Ricans, respectively. After adjusting for the confounding effects of age, sex, income, and education status, no statistically significant differences were found in the mean number of decayed teeth among the three groups of Hispanics. Puerto Rican children had an average of 2.09 filled teeth compared with an average of 1.39 and 1.43 filled teeth for Mexican Americans and Cuban Americans, respectively. In adults, Puerto Ricans and Cuban Americans had at least 40 percent higher mean number of filled teeth than Mexican Americans. Cuban American and Puerto Rican adults had about twice as many missing teeth as Mexican Americans. The pit-and-fissure tooth surfaces in children accounted for the majority of sites affected by caries.

All Hispanics had a higher prevalence of gingivitis than American adults as estimated during the 1985–86 National Institute of Dental Research (NIDR) survey of American adults. Puerto Ricans had the highest level of periodontal disease and the highest Debris Index scores among the Hispanic groups. [*Am J Public Health* 1990; 80(Suppl):66–70.]

Introduction

Dental public health practice in the United States is now facing new challenges. In addition to the political and economic changes that have affected public health programs in the last decade, there has been a significant increase in the numbers of the elderly and minorities, such as Hispanics.[1,2] As the proportion of minorities in the US increases, data on their oral health status will be needed to plan the treatment and preventive services required to meet their health needs.

Data from recent national surveys show a disturbing pattern of untreated dental caries in Blacks and Mexican Americans.[3–5] In the 1985–86 national survey conducted by the National Institute of Dental Research (NIDR),[3] Blacks had a higher mean number of decayed tooth surfaces than White non-Hispanics in all age groups between 18–19 and 60–64 years of age. The differences in mean numbers of decayed tooth surfaces between White and Black Americans ranged from 66 percent in adults 30–34 years of age to 419 percent in those 60–64 years of age. Despite the higher level of decayed teeth in Blacks, their mean number of permanent decayed and filled tooth surfaces (DFS) was lower than that of Whites, because of the lower mean number of filled teeth in Blacks. These differences between Black and White adults were similar to those found during the first National Health and Nutrition Examination Survey (NHANES I) conducted between 1971 and 1974,[6] where Blacks had a 93 percent higher mean number of decayed teeth than non-Hispanic Whites. In 1971–74, the latter had three times higher mean numbers of filled teeth than Blacks.

Information about the oral health status of Puerto Ricans, Cuban Americans, Mexican Americans, and other minorities in the United States is sparse. In a 1972–73 survey in New York City by Nikias and co-workers,[7] Puerto Ricans had a lower number of filled teeth than non-Hispanic Whites, but the mean number of filled teeth of Puerto Ricans was not different from that of Blacks. Non-Hispanic White New Yorkers had a lower mean number of decayed teeth than Blacks and Puerto Ricans. Puerto Ricans also had a higher level of gingival inflammation than non-Hispanic Whites and Blacks. Other geographically limited surveys of dental caries prevalence of Puerto Ricans and other Hispanics have reported similar findings.[8,9]

As described elsewhere in this supplement,[10] the National Center for Health Statistics (NCHS) responded to the lack of information on the health and nutritional status of Hispanics in the United States by conducting the Hispanic Health and Nutrition Examination Survey (HHANES) between 1982 and 1984. Previous reports described the prevalence of dental caries and periodontal disease of Mexican Americans living in five Southwestern states.[4,5] Dental data on Puerto Ricans and Cuban Americans became available for analysis in 1987. This report describes the prevalence of total tooth loss, dental caries, periodontal disease, and oral hygiene status in Cuban American and Puerto Rican children and adults, ages five years or older. Because of the small number of Cuban Americans and Puerto Ricans examined during HHANES compared with Mexican Americans, the age groups used in previous reports could not be used in this analysis.[4,5] Therefore, a re-analysis of the Mexican Americans' dental data was carried out in order to provide a comparison group to both Cuban Americans and Puerto Ricans.

Methods

A description of the HHANES sample design and data collected is presented elsewhere in this supplement.[10] Previous reports have summarized the criteria used during HHANES to measure the prevalence of dental caries, periodontal disease, and oral hygiene status.[4,5]

In this report, data will be presented for 5,983 Mexican Americans, 1,192 Cuban Americans, and 2,226 Puerto Ricans between the ages of five and 74 years with dental data. All those who identified themselves as Mexican Americans in Arizona, California, Colorado, New Mexico, and Texas; as Puerto Ricans in the New York City area; and as Cuban Americans in Dade County, Florida, were included in this study. Out of those medically examined, only 2.97, 1.25, and 2.36 percent of Mexican Americans, Cuban Americans, and Puerto Ricans, respectively, between six months through 74

Note: Author affiliations are listed elsewhere under CONTRIBUTORS.

years of age, did not receive a dental examination. Only the dental status of the permanent teeth is described in this paper.

All means, percentages, and regression coefficients presented were computed using sampling weights as described by Delgado and co-workers[10] and Landis, *et al.*[11] Also, adjustment for clustering effects was carried out when standard errors of the means, percentages, and regression coefficients were computed using the design effects provided by the NCHS and published in Public Use Data Tape Documentation (Dental Health, Tape number 6505). The dental variables used in the analysis were: total tooth loss; number of decayed, filled, and missing teeth (and tooth surfaces) for each individual; the Periodontal Index (PI) scores,[12] and periodontal classification based upon the PI score (no periodontal disease, gingivitis only, periodontal pockets) and Debris Index (DI) and Calculus Index (CI) scores for those with at least one tooth in their mouth.[13]

Differences in the dental status among the three Hispanics groups were considered present only if they remained after adjusting for the confounding effects of age, sex, income, and education. Multiple linear regression analysis was used during adjustment when continuous variables were compared, and logistic regression analysis was used when a dichotomous variable was the dependent variable. This approach was used because of the differences among the three groups in age, income, and education levels: the Puerto Rican sample included a higher percentage of individuals with an average family income below $10,000 than the other groups and about 43 percent of the Puerto Ricans examined were below the poverty line (computed by dividing the total household income by the total income determined necessary [from federal guidelines] to maintain a family on a nutritionally adequate diet; a ratio of 1.0 was considered as the poverty line) compared with about 29 and 22 percent of the Mexican and Cuban Americans, respectively.

All analyses were carried out using OSIRIS. IV, supported by the Institute for Survey Research, University of Michigan.

Results

Total Tooth Loss

Cuban Americans had a significantly higher prevalence of total tooth loss than both Mexican Americans and Puerto Ricans (Table 1). When the confounding effects of age, sex, income, and education status were accounted for using a logistic regression model, however, no significant differences in the prevalence of total loss among the three groups were found (data available on request to authors).

TABLE 1—Number of Mexican Americans, Cuban Americans, and Puerto Ricans, Ages Five through 74 Years, by Edentulous Status and Age, HHANES, 1982–84

	Mexican Americans		Cuban Americans		Puerto Ricans	
Age (years)	N*	%E**	N	%E	N	%E
5–17	2,528	.00	294	0.00	939	.00
18–44	2,252	.30	390	1.30	799	1.40
45–74	1,203	14.40	508	14.90	528	12.90
All Ages	5,983	2.60	1,192	6.10	2,266	2.80

*N = Total Number of Hispanics ages 5 through 74 years.

**%E = Percent edentulous (with total tooth loss) and, henceforth, were not included in Tables describing the status of individuals with teeth.

Dental Caries Prevalence

Among dentate Hispanics, ages five through 44 years, Puerto Ricans had a significantly higher mean number of decayed teeth than Cuban Americans (Table 2), but not Mexican Americans; this difference disappeared, however, when age, sex, income, and education status were accounted for using a multiple linear regression model (data available on request to authors). In older Hispanics, Mexican Americans had a significantly higher mean number of decayed teeth than either Cuban Americans or Puerto Ricans but, again the difference disappeared after accounting for the confounding effects of age, sex, income, and education status, using a multiple linear regression model (Table 3). Hispanic adults had significantly higher mean number of decayed teeth than Hispanic children.

Puerto Rican children had an average of 2.09 filled teeth compared with an average of 1.39 and 1.43 filled teeth for Mexican American and Cuban American children, respectively, a difference which remained statistically significant in the regression analysis (data available on request to authors). In adults, Puerto Ricans and Cuban Americans had at least 40 percent higher mean number of filled teeth than Mexican Americans (Table 2). In Hispanics older than 45 years, the mean number of filled teeth of Cuban Americans and Puerto Ricans was significantly higher in the regression analysis than that of Mexican Americans, with Cuban Americans having the highest mean number of filled teeth of the three groups.

Cuban American and Puerto Rican adults had approximately two times more missing teeth than Mexican Americans (Table 2). This difference did not disappear after accounting for the previously described confounding variables. Mexican American adults (18–74 years of age) who were

TABLE 2—Mean Number of Decayed, Missing, and Filled Permanent Teeth (and 95 Percent Confidence Intervals) of the Means of Dentate Mexican Americans, Cuban Americans, and Puerto Ricans by Age, HHANES, 1982–84

Age in Years	Decayed	Missing	Filled
Mexican Americans			
5–17	.70	.04	1.39
	(0.64, 0.76)	(0.03, 0.05)	(1.21, 1.57)
18–44	1.45	1.52	4.56
	(1.31, 1.59)	(1.39, 1.65)	(4.18, 4.94)
45–74	1.49	5.24	3.97
	(1.26, 1.71)	(4.85, 5.63)	(3.45, 4.49)
Total	1.19	1.56	3.36
	(1.11, 1.27)	(1.46, 1.66)	(3.15, 3.57)
Cuban Americans			
5–17	.51	.13	1.43
	(0.35, 0.66)	(0.07, 0.19)	(1.13, 1.73)
18–44	1.02	3.93	7.70
	(0.80, 1.24)	(3.46, 4.39)	(7.15, 8.24)
45–74	0.85	9.19	6.50
	(0.70, 1.00)	(8.59, 9.78)	(6.00, 6.99)
Total	0.85	4.87	5.88
	(0.74, 0.96)	(4.52, 5.21)	(5.56, 6.20)
Puerto Ricans			
5–17	.76	.11	2.09
	(0.67, 0.88)	(0.06, 0.16)	(1.80, 2.38)
18–44	1.40	3.58	7.43
	(1.27, 1.53)	(3.09, 4.07)	(6.89, 7.96)
45–74	.97	10.43	5.53
	(0.78, 1.16)	(9.25, 11.42)	(5.00, 6.06)
Total	1.10	3.26	5.15
	(1.00, 1.20)	(2.90, 3.62)	(4.83, 5.47)

TABLE 3—Multiple Linear Regression Coefficients of the Relation between Number of Decayed Teeth and Hispanic Identifier, Sex, Income, and Education: Age Group 45 to 74 Years of Age, HHANES 1982–84

Variables	Regression Coefficient	Standard Error*
Constant	2.66	.42
Hispanic Group		
Mexican American	.37	.23
Cuban American	−.04	.32
Sex (males, females)	−.48	.14
Income (Less than $10,000, $10,000–<$20,000, $20,000+)	−.16	.09
Education (Never, primary, high school, college)	−.31	.10

*Standard errors were adjusted for clustering effects (see Methods section). Confidence intervals were computed with adjustment for design effects as described in the Delgado, *et al.*, paper.[10]

below the poverty line had a significantly higher mean number of decayed teeth and a lower mean number of filled teeth than those above the poverty line. Differences in the mean numbers of decayed and filled teeth, by age groups, between Puerto Ricans and Cuban Americans below and above the poverty line were inconsistent because of small sample sizes.

Table 4 presents the distribution of dental caries by the type of tooth surface affected in Hispanic children. Puerto Ricans ages five to 17 years had significantly higher DFS (decayed and filled tooth surfaces) scores on occlusal surfaces than Mexican and Cuban Americans. In the buccal and lingual tooth surfaces of posterior teeth, Cuban American children had a significantly lower mean DFS than did Mexican Americans and Puerto Ricans. No significant differences were found among the children of the three Hispanic groups in the DFS of mesial and distal tooth surfaces of posterior or anterior teeth. Puerto Rican children had 58 percent of the carious activity occurring on occlusal surfaces, and the Cuban American and Mexican American children had 56 and 53 percent, respectively. If the buccal and lingual tooth surfaces are included, then the percent age of pits-and-fissures of posterior tooth surfaces affected by dental caries will increase to about 79, 85, and 85 percent of the total DFS scores in Cuban Americans, Puerto Ricans, and Mexican Americans, respectively.

Periodontal Disease

Puerto Ricans, ages five through 44 years of age, had a significantly higher prevalence of gingivitis than Cuban Americans and Mexican Americans (Table 5). In the oldest age group, Puerto Ricans and Mexican Americans had a significantly higher prevalence of periodontal pockets than Cuban Americans. Puerto Ricans, 18 to 74 years of age, also had significantly higher mean DI scores than Cuban Americans (Table 6). In a regression model controlling for the confounding effects of age, sex, diabetes (presence or absence of diagnosed diabetes), and oral hygiene status, Puerto Ricans remained with the highest prevalence of periodontal disease relative to the other two Hispanic groups. Diabetes status was included because it has been shown in a longitudinal study that diabetics had more gingivitis and periodontal pocketing than non-diabetics.[14]

Discussion

Untreated dental caries in Hispanic children and adults and the relatively high mean number of missing teeth in Cuban American and Puerto Rican adults compared with Mexican Americans and other Americans[3] are the two most important oral health problems facing Hispanics. It is important to note that Hispanic adults had twice higher mean number of decayed teeth than Hispanic children. This finding suggests that there is a strong cohort effect in the Hispanic population with different age groups experiencing different exposure to preventive and dental services. The presence of decayed teeth, however, does not mean that treatment of all carious lesions is urgently needed. Unfortunately, the criteria used during HHANES to measure caries[15] do not provide information about the severity of the carious lesions and therefore a decision on what mix of preventive and treatment services will be needed cannot be made. Also, the NIDR Restorative Treatment Need Index,[15] used during HHANES, does not provide information about the need for fissure sealants and whether small carious lesions should be restored or followed with recommendation for fluoride application to arrest the progress of the lesion. Given the concentration of carious activity in the pits and fissures of posterior teeth in Hispanic children, however, the effectiveness of fissure sealants in preventing dental caries on those surfaces will be higher than other preventive methods.

In a previous analysis of National Health Interview Survey (NHIS) data collected during 1978–80, Mexican Americans, Cuban Americans, and Puerto Ricans visited the dentist less frequently than White non-Hispanics.[16] Mexican Americans had the lowest proportion of persons who went to a dentist within a year compared with Puerto Ricans and Cuban Americans. Mexican American use of dental services resembled the pattern of dental care received by Black

TABLE 4—Mean Number of Decayed and Filled Tooth Surfaces (DFS) by Type of Tooth Surface Affected in Dentate Mexican Americans, Cuban Americans, and Puerto-Rican Children, 5–17 Years of Age, HHANES, 1982–84

	Posterior Teeth			Anterior Teeth	
	Occlusal	BL	MD	BL	MD
Mexican Americans	1.79 (1.64, 1.94)	1.06 (0.97, 1.15)	.30 (0.24, 0.36)	.10 (0.07, 0.13)	.10 (0.06, 0.14)
Cuban Americans	1.67 (1.38, 1.96)	.66 (0.51, 0.81)	.38 (0.23, 0.53)	.09 (0.03, 0.15)	.14 (0.04, 0.24)
Puerto Ricans	2.56 (2.28, 2.83)	1.29 (1.12, 1.46)	.36 (0.25, 0.47)	.06 (0.03, 0.09)	.10 (0.05, 0.15)

Occ = Occlusal
BL = Buccal, Lingual
MD = Mesial, Distal

TABLE 5—Percent Distribution of Dentate Mexican Americans, Cuban Americans, and Puerto Ricans by Periodontal Disease Status* and Age, HHANES, 1982–84

	Gingivitis Only			Periodontitis		
Age (years)	Mexican American (95% CI)	Cuban American (95% CI)	Puerto Rican (95% CI)	Mexican American (95% CI)	Cuban American (95% CI)	Puerto Rican (95% CI)
5–17	76.6	65.1	90.0	.2	.0	.1
	(73.4, 79.8)	(59.4, 70.8)	(87.3, 92.7)	(−.14, .54)	—	(−.2, .4)
18–44	82.4	81.2	90.0	7.0	2.1	8.2
	(79.3, 85.5)	(77.1, 85.3)	(87.1, 92.9)	(4.9, 9.1)	(.6, 3.6)	(5.5, 10.9)
45–74	59.1	72.3	65.1	35.7	20.8	32.4
	(53.2, 65.0)	(67.9, 76.7)	(58.9, 71.3)	(29.9, 41.5)	(16.8, 24.8)	(26.3, 38.5)

*Periodontal Disease Status is based on Periodontal Index Scores.[12] Those with no periodontal disease are not included and therefore the percentages do not add up to 100.

TABLE 6—Mean Number of Debris Index (DI) and Calculus Index (CI) Scores[13] (and 95 Percent Confidence Intervals) of the Means for Dentate Mexican Americans, Cuban Americans, and Puerto Ricans, Ages Five through 74 Years, HHANES, 1982–84

Age (years)	Mexican American	Cuban American	Puerto Rican
Debris Index scores			
5–17	.80	.72	.78
	(0.76, 0.84)	(0.66, 0.78)	(0.74, 0.82)
18–44	.75	.62	.86
	(0.70, 0.80)	(0.57, 0.67)	(0.81, 0.91)
45–74	.94	.72	1.04
	(0.84, 1.04)	(0.65, 0.79)	(0.95, 1.13)
Calculus Index scores			
5–17	.09	.04	.16
	(0.08, 0.10)	(0.02, 0.06)	(0.13, 0.19)
18–44	.67	.43	.77
	(0.63, 0.71)	(0.37, 0.49)	(0.70, 0.84)
45–74	1.30	.75	1.23
	(1.22, 1.38)	(0.68, 0.82)	(1.11, 1.35)

Americans. Puerto Ricans were the only group for whom dental visits were not significantly influenced by family income level. In this study, the dental caries prevalence of Mexican Americans parallels that found in Black Americans during the 1985–86 NIDR survey.[3] (Comparisons between HHANES findings and other national surveys are limited by the differences in sampling strategies. Examination criteria used during HHANES, however, were those used by NIDR, and the training of the dental examiners in HHANES was also carried out by NIDR staff.) Puerto Ricans and Cuban Americans had a dental caries distribution that was in between those of Mexican Americans and Blacks, and White Americans in the NIDR survey. The influence of dental visit patterns, income, and dental education may partly explain the differences in the mean number of missing teeth among the three Hispanic groups.

The results of the analysis of HHANES dental data reported here raise a number of methodological questions about the meaning of the DMF score. Conventionally, the DMF score has always been analyzed as a single composite index. Moreover, analysis of the DMF score of an individual assumed that all filled teeth have been treated because they were previously carious. The filled component of the DMF index represents a decision made previously by the dentist of the examinee. As has been documented by several investigators[17,18] the diagnostic criteria used by dentists to decide on need for restorative treatment are plagued with considerable variations. In a study in Scotland, where 15 general practitioners examined 18 patients and estimated their restorative treatment needs, the range of teeth needing treatment for each patient varied according to the examining dentist.[17] For the 18 patients together, the costs of the treatment planned by the 15 dentists ranged between 92.5 and 478.6 British pounds. These findings raise doubts about the validity of analytic methods for the DMF index that combine the three components of the index as a single variable.

As in Mexican Americans, the prevalence of gingival inflammation and periodontitis in Puerto Ricans and Cuban Americans is higher than national averages.[3] Puerto Ricans had the highest prevalence of gingivitis. Puerto Ricans and Mexican Americans also had a higher prevalence of calculus and debris than Cuban Americans. The level of DI and CI scores in Mexican Americans and Puerto Ricans, 18–74 years of age, is higher than that determined to be compatible with an acceptably low level of periodontal disease.[19]

In conclusion, the oral health status of Hispanic-Americans can be summarized as follows:

- Total tooth loss is not a public health problem in Hispanic Americans examined during HHANES. However, the dentate Cuban American and Puerto Rican adults have about twice as many missing teeth than White non-Hispanics.
- In children, while the prevalence of dental caries is low, using the World Health Organization criteria,[20] the tooth surfaces most affected by dental caries are the pits and fissures of posterior teeth rather than smooth tooth surfaces. These findings support the need for sealant programs to prevent dental caries.
- Hispanic adults had twice higher mean number of decayed teeth than Hispanic children. Puerto Ricans and Cuban Americans had a higher mean number of filled and missing teeth than Mexican Americans. The distribution of dental caries in Mexican Americans was similar to that found in Blacks in 1985–86.[3] Puerto Ricans and Cuban Americans had a dental caries distribution between that of Black and White Americans.[3]
- The prevalence of gingivitis in Hispanics is higher than that found in White non-Hispanics.[3] Puerto Rican children and adults had the highest prevalence of gingivitis among the Hispanic groups.

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Public Health Policy Forum

Disease Control Priorities in Developing Countries: Health Policy Responses to Epidemiological Change

Dean T. Jamison, PhD, and W. Henry Mosley, MD, MPH

ABSTRACT

Health systems in developing countries are facing major challenges in the 1990s and beyond because of a growing epidemiological diversity as a consequence of rapid economic development and declining fertility. The infectious and parasitic diseases of childhood must remain a priority at the same time the chronic diseases among adults are emerging as a serious problem. Health policymakers must engage in undertaking an epidemiological and economic analysis of the major disease problems, evaluating the cost-effectiveness of alternative intervention strategies; designing health care delivery systems; and, choosing what governments can do through persuasion, taxation, regulation, and provision of services. The World Bank has commissioned studies of over two dozen diseases in developing countries which have confirmed the priority of child survival interventions and revealed that interventions for many neglected and emerging adult health problems have comparable cost-effectiveness. Most developing countries lack information about most major diseases among adults, reflecting lack of national capacities in epidemiological and economic analyses, health technology assessment, and environmental monitoring and control. There is a critical need for national and international investment in capacity building and essential national health research to build the base for health policies. (*Am J Public Health* 1991;81:15–22)

Introduction

The policy debate in international health has often been polarized around conflicting viewpoints on such issues as preventive versus curative services, selective versus comprehensive primary health care, or integrated versus vertical programs. As we approach the 21st century, it is becoming clear that framing the issues in these terms will not enlighten the policy process, primarily because it limits the options largely to actions that can be carried out directly by ministries of health. Profound social and economic transformations are projected to impact on health in the developing countries in the 1990s and beyond; implications for the epidemiological profiles of these countries will be dramatic.[2–4] A more comprehensive analytical approach is required to formulate health policies that will not only respond to but actually guide the development process to maximize its health gains, minimize its potential adverse consequences, and deal cost-effectively with the emerging quantitative importance of noncommunicable diseases.

Over the next decade there will indeed remain perhaps 30 to 40 countries in the lowest income bracket where the health problems will be dominated by the infectious and parasitic diseases of childhood along with undernutrition and high fertility.[5] Most of these are in Sub-Saharan Africa and South Asia, and intervention strategies must continue to focus on establishing basic infrastructures and delivering established, low-cost technologies.[6] At the same time, however, there will be a larger group of countries, predominantly in Latin America and East Asia, that are newly emerging into the middle income category. Most of these countries are not free of the infectious diseases, yet at the same time are beginning to face a new set of health problems related to rapid urbanization and industrialization.[3,7,8] These include: injuries, occupational diseases and, increasingly, preventable chronic diseases among an aging population.[9] Such a growing epidemiological diversity among and within countries in the developing world demands a flexible approach to international health policy formulation.

Policy Development

It is useful to consider health policy development as involving three complementary tasks: first, *identifying the major disease problems,* assessing their social and economic consequences, and evaluating the costs and effectiveness of alternative intervention strategies; second, *designing health care delivery systems* including establishing the human and physical infrastructures, providing for drugs and logistical support, and developing managerial capacities and funding mechanisms; and third, *defining and choosing what governments can do* through the full range of policy instruments that are at their disposal in the areas

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Address reprint requests to either author. Dr. Jamison is with the Department of Community Health Sciences, University of California, Los Angeles, CA 90024. Dr. Mosley is with the Department of Population Dynamics, Johns Hopkins University, School of Hygiene and Public Health, Baltimore, MD 21205.

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of persuasion, taxation, regulation and the provision of services.

Like a three-legged stool, coherent health policies require a solid foundation of technical analysis in all three areas in order to identify effective strategies and establish sustainable programs. Through the 1960s and 1970s the second task, designing institution-based delivery systems, (generally following a Western medical model) was the dominant theme in health policy formulation for the developing world. It was in this period that the policy debates centered around the relative allocation of resources for preventive versus curative services and, when disease eradication programs were being implemented, vertical versus integrated programs. The Alma Ata Conference on Primary Health Care in 1978 introduced the principle of "Health For All."[10] Conceptually this encompassed all health problems in the population, was community-based, and involved all sectors of government. Given the resource constraints in the health sector, many health professionals and donors narrowed the agenda, limiting consideration to childhood communicable diseases.[11] This led to selective primary health care programs focusing largely on population-based technical interventions such as oral rehydration therapy and immunizations for child survival[12] which, in turn, led to policy debates about the relative merits of selective versus comprehensive primary health care interventions.[13]

The third task, assessing the full array of government instruments available, particularly those outside the traditional purview of ministries of health (such as regulation and taxation) to protect and promote the health of populations is not well established in much of the developing world. The full range of instruments is, however, fully entering the public debate on policy in a number of industrialized countries. For example, in the United States the process was initiated with the publication of "Healthy People: The Surgeon General's Report on Health Promotion and Disease Prevention" in 1979 which defined a set of health goals for 1990 for infants, children, adolescents, adults, and the aged.[14] To achieve these goals, 15 priority areas for intervention were identified and grouped under three strategic approaches: preventive health services, health protection, and health promotion.[15] Preventive health services included: primary preventive interventions such as immunizations and family planning, secondary prevention for conditions like high blood pressure as well as curative services for such conditions as sexually transmitted diseases. Health protection encompassed areas where regulation and taxation would be important including toxic agent control, occupational safety, accident prevention, and fluoridation. The health promotion strategies—involving behavioral changes related to smoking, alcohol and drug abuse, diet and exercise—required government interventions not only in education but also through regulation and taxation.

The process of setting national objectives in the US has now been extended with *Healthy People 2000: National Health Promotion and Disease Prevention Objectives.*[16] Based on three years consultation with more than 300 national organizations and state health departments, the goals have been reformulated to go beyond mortality reduction: to increase the span of healthy, disability-free life; to reduce disparities in health by giving more attention to the vulnerable high risk groups such as the poor, minorities, and the elderly; and, to increase access to basic health services.[16] In this context, new priority areas have been added such as cancer screening, AIDS, preventive health services, and surveillance.

Three factors have been key to the successful mobilization of public and professional interest and debate in the United States.[17] First, there was a clear definition of the health problems of concern and their social and economic as well as medical determinants. Second, the 15 priority areas selected for health improvement were further subdivided into 227 health promotion objectives for the 1980s.[18] These detailed objectives represented precise, quantifiable management strategies which were realistic in terms of resource constraints and time frame for implementation, yet demanding enough to push the system to new levels of capability. Third, the *process* of problem identification and health policy formulation involved the broadest participation of all segments of society.[19] On the one hand, a considerable breadth of expertise was drawn from the fields of epidemiology, management sciences, economics, behavioral sciences, and communications to define the problems and issues to be addressed and propose strategic alternatives. On the other hand, there was broad political and public participation in devising solutions. The necessity of this third step (which is rare in developing countries) cannot be overemphasized: the policy formulation process should be an educational experience about the health problems for all concerned since it will be necessary to reach a wide national consensus on issues, policies and implementation strategies which will cut across many sectors including finance, trade, agriculture, transportation, industry, education, and others. It must also be emphasized that successful mobilization of debate and concern is far from sufficient, as the example of the US also, unfortunately, attests. Where the political will for public action is lacking and resources are unallocated, health outcome indicators, such as infant mortality rates among minority groups, can remain shamefully high.

A first task in a comprehensive analysis of the health sector requires an estimation of the burden of disease in the population. This should be done in a framework that considers not only the capacity of the health system to deal with these diseases but also the rapidly changing social, economic, and political context that may be influencing both the structure and organization of the health system as well as disease risks and health behaviors in the population. Because of the extraordinary diversity among developing countries, it is unrealistic and, indeed, inappropriate to come up with global policies; rather, what is presented here is a conceptual and analytical approach with illustrations of the range of factors which must be considered in any particular setting. This approach takes into account that there will be a continuing evolution of disease control priorities based on the social and economic transformations countries are undergoing with their accompanying demographic and epidemiologic transitions.[20]

The Changing Determinants of Disease

A first determinant of the changing health picture developing countries will be experiencing is the demographic transition, largely driven by the rate of fertility decline. This will produce major changes in the age structure of populations which will have a large impact on the health picture.[21] Table 1 taken from World Bank projections for the period 1985 to 2015 illustrates the changing demographic and health picture for developing countries worldwide.[22] Overall, slower growth of the childhood population coupled with declining death rates are projected to reduce the health problems of all age groups. But this is more than counterbalanced by the

TABLE 1—Health Problems of Different Age Groups

Age Group	Population in Developing Countries, millions 1985	Population in Developing Countries, millions 2015	Deaths in Developing Countries, millions 1985	Deaths in Developing Countries, millions 2015	Important Health Problems: Problems on the Unfinished Agenda	Important Health Problems: Neglected and Emerging Problems
Young children (0–4 years)	490	626	14.6	7.5	Diarrheal disease ARI Measles, tetanus, polio Micronutrient deficiencies Malaria	Injury
School-age Children (5–14 years)	885	1196	1.6	1.3	Schistosomiasis Geohelminth infection	Adolescent pregnancy Disability
Young adults (15–44 years)	1667	2918	5.0	6.0	Maternal mortality Malaria Excess fertility	Tuberculosis Injury AIDS STDs Mental Illness
Middle-aged (45–64 years)	474	1131	5.9	10.4		Cardiovascular disease Cancers COPD
Elderly (65+ years)	153	358	11.0	22.5		Disability Depression
TOTALS	3669	6229	37.9	47.7		

SOURCE: Jamison and Mosley, (forthcoming)[1]

NOTE: Many of the conditions for older age groups manifest themselves clinically long after the processes leading to the clinical condition have been initiated; preventive intervention will therefore need to be directed to younger ages.

AIDS: acquired immunodeficiency syndrome
STDs: sexually transmitted diseases
COPD: chronic obstructive pulmonary disease
ARI: acute respiratory infection

rapid increase in the adult and aging population. Regionally the picture is much more diverse. The population of children under age five in Latin America and Asia will increase by only 2 percent to 5 percent, respectively, in size. By contrast, this age group is projected to grow 38 percent in the Middle East and 70 percent in Sub-Saharan Africa from 1985 to 2015. Clearly the future requirements for maternal and child health services will be quite different across these regions. At the other end of the age spectrum a very different picture emerges. In all of these regions the populations over age 65 will be dramatically growing in size with projected increases ranging from 134 percent to 164 percent. This continuing growth in the older age groups is due to a phenomenon called the "momentum" of population growth; it is a consequence of the large numbers of children born in the past when fertility was high who advance through older age groups over time. The result is that essentially all developing countries will have to provide for the health needs of a rapidly increasing adult and aged population for many years into the future.[4] A broad assessment of how to meet these needs may be found in Feachem, *et al*, (forthcoming).[23] At the same time, in those countries where the fertility transition is barely underway (largely in Sub-Saharan Africa), policymakers must simultaneously plan to continue to rapidly expand maternal and child health services.

A second determinant of the changing health picture is the major social and economic trends which are transforming the risk factors themselves.[24] The most obvious indicator of these trends is the shift from rural to urban living. In 1985, barely over 30 percent of the population of the less developed regions lived in urban areas, but this pattern is changing rapidly.[25] The urban population will reach 40 percent by the year 2000 and surpass 50 percent by the year 2015. Already seven of the world's 12 megacities (population 10-million or more) are in developing countries. By the year 2000 the developing world will add six more cities to this number.

A shift from a rural subsistence economy to an urban market-oriented industrial economy is generally associated with reductions in risks to communicable diseases because of better sanitation in urban areas. At the same time, however, economic growth brings with it new health problems.[26] Very high rates of injuries related to motor vehicles, industrial accidents, and toxic chemicals (e.g., pesticides) are usually seen in developing countries because of a lack of resources and institutions to establish and enforce safety measures.[27] Undernutrition may diminish because of improved markets, only to be replaced by overnutrition with rising risks of death due to obesity, hypertension, atherosclerosis, and diabetes.[28] Rising incomes also bring changes in lifestyle including increases in smoking, alcohol use, and substance abuse, all of which are expected to increase the risk for chronic diseases.

An analysis of the tobacco-related diseases provides an excellent illustration of how a disease-oriented approach can inform health policy.[29] In the United States it is now well established that tobacco use is responsible for more than 30 percent of all cancer deaths as well as being among the strongest risk factors for chronic obstructive pulmonary disease and ischemic heart disease.[30,31] In recent years, rising rates of lung cancer associated with cigarette consumption have been observed in Japan, Singapore, and Shanghai.[32] A recent analysis by the World Bank has documented a direct relationship between tobacco consumption and higher levels of national income among developing countries.[33] This relationship is not coincidental. Rather, it is driven by promotion, pricing, tax policies, and even international trade relations. The impact of the latter is best attested by the rapid increases in cigarette consumption among young people in several east Asian countries where multinational tobacco corporations have recently opened new markets, assisted in some cases by aggressive (even coercive) trade policies of the United States Government.[34]

Taking China as an example, Peto projects that if current smoking trends continue due to a failure to introduce the array of policies necessary to reduce tobacco consumption in the next several years, an estimated two million Chinese men will die annually from tobacco-related health problems by the year 2025.[35] Cumulatively, one may anticipate 50 million tobacco-related deaths among the 500 million Chinese now under the age of 20. National policies to protect populations from tobacco-related diseases clearly must move beyond the usual realm of the health sector into areas of regulation and taxation of the industrial enterprise related to the manufacture and distribution of tobacco products and, in many countries including the United States, ultimately into a restructuring of rural economies where tobacco is produced.[36-38]

The AIDS (acquired immunodeficiency syndrome) epidemic provides another illustration of some of the links between economic development and disease transmission. In assessing the situation in Sub-Saharan Africa, Over and Piot document a strong correlation between the rate of HIV infection in 18 cities, a high ratio of males to females in urban centers and a low level of female education.[39] The implications of these findings are that AIDS transmission is facilitated by development strategies which favor males for urban employment. This generates a high demand for prostitutes; prostitution, in turn, is facilitated by low levels of female education so that women have few alternative economic opportunities open to them. An AIDS-control program must go beyond promoting the use of condoms and treating sexually transmitted diseases (STDs) and consider broader social policies relating to female education and labor force participation.

Epidemiologic Polarization

The epidemiologic transition in the developing world is neither a steady nor a uniform process. The great variations in life expectancy between countries attests to this; as of 1988, a dozen Sub-Saharan African countries still had life expectancies under 50 years, while an equal number of countries in Latin America and Asia (including China) had life expectancies of 70 years or greater.[40] Additionally, within developing countries (and developed countries as well) wide disparities in health conditions across different social classes or different regions are not uncommon. Frenk and colleagues have referred to these disparities as "epidemiologic polarization" and "epidemiological stagnation."[3,41] These disparities are not limited to the infectious diseases of infants and children but include the chronic diseases of adults as well.[7,23] A centrally important conclusion for health policy is that in most developing countries pre- and post-epidemiological transition problems will coexist.[41] As Foege and Henderson have observed, the developing countries ". . .will not have the luxury of dealing with two kinds of problems sequentially. For the remainder of this century they will be dealing with both simultaneously."[42]

The increasing burden of chronic disease is initially likely to affect the relatively more affluent and politically vocal older groups who are growing in numbers. This being the case, governments will need to take great care to assure that the infectious diseases which predominantly affect children and the poor are not neglected in the face of resource demands placed in large measure by the more affluent. The challenge is to develop equitable policies and strategies to meet these unprecedented conditions in the developing world.

Assessing Disease Control Priorities

The World Bank, as a part of an exercise to assess the cost and effectiveness of alternative intervention strategies for over two dozen acute and chronic diseases of adults and children, commissioned a series of analytical studies, each co-authored by individuals combining economic, epidemiological, and clinical expertise. Cost effectiveness was expressed in terms of dollar cost per discounted healthy life years gained (DHLY). The details of these studies are reported elsewhere.[43] Several general conclusions which emerged from this undertaking will be described here in terms of their implications for the process of policy development in developing countries.

First, while child survival interventions (immunizations, provision of antenatal and delivery care, vitamin A supplementation, improvements in domestic hygiene, oral rehydration therapy [ORT] for diarrheal diseases in high mortality environments, and antibiotic therapy for acute respiratory infections) appear to be highly cost-effective ($5-50 per DHLY), there were a range of adult health interventions which were estimated to be equally cost-effective. Among the most prominent are: anti-smoking campaigns plus tobacco taxes, passive case finding and short-course chemotherapy for tuberculosis, the targeted management of sexually transmitted diseases, cataract surgery, the use of condoms to prevent HIV transmission, and hepatitis B immunization to prevent liver cancer and cirrhosis. At the other extreme, the cost per DHLY for a number of adult diseases is so high (greater than $1000) that public policy should probably discourage their use in settings where health resources are severely constrained. Among these are coronary bypass surgery, mitral valve replacement for rheumatic heart disease, medical management for hypertension, and tertiary management of lung, liver, esophageal, and stomach cancer. With reference to cancer, it should be noted that cancer pain control appears to be a relatively cost-effective intervention for adults ($150 per DHLY) reaching the same level as: maternity care services for maternal mortality; public preventive campaigns to prevent cardiovascular disease; medical management of angina; and insulin management of insulin-dependent diabetes melitis.

This analytical approach revealed a number of neglected and emerging health problems which should be accorded far higher priority. Topping the list of emerging problems are the tobacco-related diseases which, interestingly enough, were not even mentioned in the 1980 World Bank Health Sector Report.[44] Tuberculosis appears to be the greatest "neglected" disease.[45] In 1990 there will be an estimated 8.4 million new cases and 2.7 million deaths, over two-thirds of which will be among productive adults (ages 15–59), primarily the poor. Significantly, this disease alone accounts for about 37 percent of an estimated seven million avoidable adult deaths in the developing world.

Some findings challenge the current international donor commitment to global priority setting. Provision of ORT *in low mortality environments* (at $200 per DHLY) is estimated to be about 20 times less cost effective than tuberculosis control (using passive case detection and short course chemotheraphy) or leprosy rehabilitation. Even cervical cancer screening (Papanicolaou smears) are estimated to be an equally good buy for the money. In *high mortality environments,* in contrast, ORT can be quite cost-effective, underscoring the need for analysis at the country or even district level.

A major observation from the World Bank exercise is the paucity of empirical data from a majority of the developing countries about most of the major chronic diseases of adults.[23,46] Considering that, on a global level, more than half the world's deaths due to cancers and cardiovascular diseases, three-fourths of the injury deaths, and 85 percent of the chronic obstructive pulmonary disease deaths occur in developing countries, the virtual absence of any information about the levels, trends, and determinants of these conditions for the majority of the world's population is truly tragic. Given this lack of developing country data, most of the estimates of the cost-effectiveness of interventions for the World Bank study had to be derived from research carried out in the industrialized world. While these estimates are probably reasonable in grossly illustrating the relative cost-effectiveness of alternative interventions, the actual cost-effectiveness of a particular intervention could easily vary by two- to tenfold or more in any given situation, depending on a host of local factors. For example, with interventions, like immunizations, that require total population coverage, cost-effectiveness is strongly conditioned by the underlying level of disease incidence and case fatality rates. For therapeutic interventions, cost-effectiveness is determined not only by the institutional costs of reaching the target group and then making the correct diagnosis, but also by the probability that appropriate therapy will be prescribed by the caregiver and that the patient will actually comply with the treatment regimen.[47–50] Similarly, the costs of monitoring and enforcing regulatory approaches depend upon the reach and capability of the governmental infrastructure.

Implications for National Governments

It should be clear that there is an urgent need to reassess health sector priorities in developing countries. National health policy assessment can be usefully informed with a consideration of global trends and findings such as presented above, but it must quickly evolve to a critical analysis of the local epidemiological picture cast in its social-cultural, administrative, economic, and political context. Such an approach to policy formulation must begin by assessing the full range of health problems in a society; examining broadly their demographic, social, and economic determinants; and identifying alternative courses of action in all of these areas which could improve the situation. This last step must take into consideration the costs and effectiveness of all possible instruments of government intervention including taxation and regulation as well as persuasion and the direct provision of services.[4]

Many governments will need to create new institutions and reconfigure old ones to build the capacities required for policy analysis and implementation. Among the analytical and technical capacities required are the following:

- *Demographic analysis*—These capabilities provide the fundamental underpinnings for a population-based health system. Demographic data provide the basis for designing intervention strategies as well as for assessing the impact of the disease burden on the population.
- *Epidemiological surveillance*—While some epidemiological surveillance is carried out in most developing countries, it is usually limited to selected infectious diseases. These capacities will need to be greatly strengthened as health program strategies move more toward interventions involving regulation, taxation, subsidies, and information programs in order to reduce acute and chronic disease risks by changing behaviors and improving environmental conditions.
- *Economic and financial analysis*—These capabilities will be essential to measure the cost-effectiveness of alternative intervention strategies as well as to assess the overall claim of the health sector on scarce development resources.
- *Health technology assessment and control*—Institutional capabilities in this area must include not only the assessment of the effectiveness of new drugs, vaccines or treatments, but also their costs and benefits when introduced into the health system.
- *Environmental monitoring and control*—In most developing countries capabilities in this area must be greatly expanded beyond the traditional water and sanitation activities to monitor and control (through taxation as well as regulation) a much broader range of environmental risks including air pollution, toxic wastes, traffic hazards, and other injury risks.
- *Occupational safety*—Rapid urbanization and industrialization in developing countries is often associated with risks of work-related injury five to 10 times higher than in developed countries. There is a major need to develop the capacity to monitor and control occupational health risks.

While many of these institutional capacities exist in some measure in health ministries, their functions are limited for a variety of reasons including lack of professionally qualified personnel, limited resources and, particularly, lack of enforceable statutory authority. If health systems are to be transformed to meet the challenges of the future, a high priority must be given to empowering health ministries to carry out the monitoring and regulatory tasks needed to effectively function in new areas. In addition, mechanisms need to be established that will give them an appropriate voice in tax policy. Ultimately these capacities will need to be developed at regional and local levels within countries as well as at the national level.

Implications for International Aid

Official development financing for health by bilateral and multilateral donors in 1986 amounted to $3.7 billion with the US contribution accounting for 16 percent of the total.[51] While external support represents only a small fraction of the total public and private sector health expenditures in developing countries, in fact, because much of the donor assistance is directed toward the poorest countries, its proportional impact in influencing health policies and program strategies is relatively large. This is particularly true when it promotes policies which lead to a reallocation of national resources. The impact of the donor-driven child survival initiative in promoting immunizations and oral rehydration therapy worldwide is an excellent illustration of this.[52]

TABLE 2—Instruments of Aid

Objectives	Modality of Assistance: Program Implementation	Modality of Assistance: Capacity Strengthening
1. Service Delivery	Supports acquisition of drugs, equipment, and technical assistance, for delivery of EPI, vector control programs, hospital services, etc.	Invests in institutional development and staff training to improve efficacy of service delivery—e.g. through improved logistics and supply systems.
2. Policy Improvement	Identifies specific areas of policy improvement (e.g. tobacco advertising bans or introduction of cost-recovery mechanisms) and includes them (usually with conditionality) as part of an assistance package.	Invests in development of policy and planning departments in ministries or universities; invests in staff training and advanced education.
3. Undertaking Research (including epidemiologic, evaluational and economic analyses)	Conducts (perhaps with involvement of aid agency or expatriate staff) research or analyses to strengthen formulation of policy or delivery of service.	Invests in national (and international) capacity for undertaking research relevant to epidemiological and economic conditions of developing countries. This capacity strengthening involves both institutional and human resource development.

In looking to the future, it is useful to categorize international aid on the basis of whether its objective is to: 1) support the provision of services; 2) influence the policy environment; or 3) support research. Table 2 divides approaches to meeting these three objectives into two modalities of assistance: program implementation and capacity strengthening. Program implementation is largely oriented toward short-term results, while capacity strengthening generally requires long-term investments in institutional development. Traditionally, the dominant role of international assistance has been to provide services and related commodities where they were unavailable or inadequate; most donor support to child survival programs falls in this category. Recently, more attention is being given to structuring aid to change the policy environment, either directly, for example, by channeling more support to private and non-profit NGO (nongovernmental organizations) sectors, or indirectly by attaching conditions to aid packages, for example, by requiring cost recovery or extension of services to underserved populations. Presently only a small fraction of aid is directed toward research, much of which, in fact, is in specialized programs such as the Tropical Disease Research Program or in a few institutions like the International Center for Diarrheal Disease Research, Bangladesh.[53]

The earlier discussion has highlighted the fact that developing countries are now moving through the epidemiologic transition at very different paces and often incompletely. As a result there is a great diversity both in health conditions and in institutional capacities among and within countries.[2] This creates a far more diffuse planning environment than existed when most of these countries were still in a pre-transitional stage two or three decades earlier. No longer can a few strategic objectives command central importance for health policy. Rather, a much broader range of relatively less prominent health conditions and associated interventions will compete for resources and attention. In light of this, and given the fact that extremely little is known about the levels, trends, and determinants of the majority of diseases afflicting populations in the developing world, the conclusion is inescapable that as far as donor assistance is concerned, far higher priority should be given to research, both building research capacity and conducting research to strengthen the formulation of health policy. This, in fact, is the conclusion reached by the Commission on Health Research for Development in its recent report.[53] The Commission specifically placed a high priority on local epidemiological and operational analyses which they labeled as "Essential National Health Research" (ENHR).

Broadening the array of diseases for consideration to encompass all ages and expanding the arena of policy options to include the full range of instruments for government intervention will demand a diversity of analytical skills. National governments will require broad technical competence not only in the biomedical and clinical sciences including epidemiology but also in economics, sociology, anthropology, the management sciences, and communication. In most countries, new institutions may need to be developed. The international donor community generally, and the United States in particular, could most effectively use its limited resources in the areas of education, training, research, and technical assistance, where it has strong comparative advantage.[54] Furthermore, because most of the major disease problems are actually of global concern (e.g. tobacco-related diseases, substance abuse, AIDS, injuries, occupational diseases, cardiovascular disease, cancer, malnutrition) much could be gained through international collaborative endeavors in these disease control efforts.

Conclusion

For over a decade the international public health community has given a priority to the communicable childhood diseases. This had been appropriate. The problems are major, the technological and epidemiological tools have become powerful, and the payoff for adapting and applying what is known is high. While the implementation strategy for selective primary health care as originally proposed by Walsh and Warren may have been oversimplified and too narrowly concerned with infectious diseases, the problem-oriented, population-based conceptual approach is fundamentally correct.[11,55] Furthermore, as more and more governments have seriously addressed the problem of child survival, the range of interventions has broadened.[56] For example, diarrhea control programs have moved from simply relying on packets of oral rehydration salts to the use of homemade fluids, the dietary management of diarrhea, the promotion of personal hygiene, and use of soap. More importantly, policy recommendations supporting child survival have now expanded widely, encompassing such issues as female education and the status of women.

The rapid demographic transition that many developing countries are undergoing is now producing great epidemiological diversity among developing countries; this demands that health policymakers consider a broader range of options. Over the next 25 years it is not unreasonable to project a 50 percent to 60 percent decline in the numbers of deaths among infants and children related to advances in child health coupled with rapidly declining birth rates which will stabilize the numbers of newborns from year to year over much of the world. Overall, however, there will be a rise in the total number of deaths as the population evolves toward an older age structure. The epidemiological transition is not occurring uniformly, however, and most countries will have a large lingering burden of communicable diseases among children, particularly in Sub-Saharan Africa where the fertility transition has hardly begun. This situation will impose the greatest challenge on health systems operating under constrained resources to select the most effective and efficient means of disease prevention, case management, and rehabilitation.

It must be emphasized that extending the range of choices means not only choosing to implement new initiatives but also choosing to reduce investments of public funds in health care activities that are clearly not cost-effective. While this may be difficult, given the traditions of Western medical practice, the fundamental issues of equity as well as efficiency are involved. Efficient resource allocation is desirable in every sector. In the health sector, however, in a very direct way, unnecessary death, disability, and illness occurs if resources are committed to one intervention when another has higher health gains per unit of expenditure. The human cost of uneconomic resource allocation is very real and very large. Perhaps even more than in other sectors, then, the imperative exists for policymakers in the health sector to undertake constant assessment of intervention cost-effectiveness.

Probably the most important advance in policy formulation that could come from a disease-control strategic approach is a willingness to consider a range of options available to governments beyond the direct provision of health services. Particularly important as countries are moving through an accelerated process of economic development is the judicious use of regulation, legislation, taxation and subsidies to promote or discourage enterprises, activities, or behaviors which may have health consequences. Also, the importance of mass communication for social mobilization to promote health cannot be over-emphasized. A high priority problem demanding such multi-faceted policy intervention and global cooperation is tobacco consumption.

It is universally recognized that improved health in the broadest sense is a fundamental indicator of the development process. As the Nobel Laureate T. W. Schultz observed, ". . .the wealth of nations (has) come to be predominantly the acquired abilities of people—their education, experience, skills, and health."[57] The task of health policy is to define clear, realistic, and measurable objectives which can guide the development process toward effectively and efficiently producing good health for all. ☐

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Patterns of Medical Employment: A Survey of Imbalances in Urban Mexico

Julio Frenk, MD, MPH, PhD, Javier Alagon, DPhil, Gustavo Nigenda, MA, Alejandro Muñoz-delRio, BSc, Cecilia Robledo, BA, Luis A. Vaq́uez-Segovia, BA, and Catalin Ramírez-Cuadra, BA

ABSTRACT

This article quantifies the magnitude and correlates of the major imbalances affecting the employment of physicians in the urban areas of Mexico. Since the early 1970s the country has experienced a rapid increase in the supply of doctors, which its health system was unable to absorb fully. In 1986, we conducted a survey in the 16 most important cities based on a probability sample of households where someone with an MD degree lived. A total of 604 physicians were interviewed for a response rate of 97 percent.

The unemployment rate was 7 percent of potentially active physicians; 11 percent held a nonmedical job, and another 11 percent exhibited low productivity and/or income. All in all, we project that 23,500 physicians in these cities were either unemployed or underemployed.

This medical employment pattern was analyzed against five independent variables: generation (i.e. the year in which the physician started medical school), gender, social origin, medical school quality, and specialty. Apart from generation, type of specialty exhibited the strongest correlation with the employment situation of a physician.

The results suggest that higher education and health care in Mexico may be producing rather than correcting social inequalities. Policy alternatives are discussed to restore a balance between the training of physicians, their gainful employment, and the health needs of the population. (*Am J Public Health* 1991; 81:23–29)

Introduction

Among the many paradoxes of contemporary health systems probably none is so indicative of failure as the coexistence of underemployed physicians with underserved populations. This is the situation that today prevails in many countries of the world.[1] For reasons that are not yet fully understood, a wide variety of nations, with very different levels of economic development and political structures,[2] experienced during the last decades substantial increases in the supply of physicians, which the health system was often unable to absorb.[3] The result has been the appearance of manpower imbalances that have assumed manifold manifestations.[4]

In order to contribute to our understanding of the dynamics of medical employment, the present article analyzes the case of physicians in the urban areas of Mexico. Apart from the intrinsic interest of the subject, the Mexican health care system offers the analytical advantage of encompassing virtually all organizational forms of medical work, from independent practice to salaried jobs in the public sector; extreme forms of imbalances such as underemployment and unemployment exist The conceptual framework and specific methodology we have developed may be relevant to other nations.

Overview: The Medical Labor Market in Mexico

From the establishment of the first course of medicine in 1578 until the beginning of the 1970s, Mexico suffered from a chronic shortage of physicians. Starting in 1967, however, the medical education system experienced an explosive expansion. Student enrollment peaked in 1980 at 93,365, compared to 28,731 in 1970;[5] the number of medical schools reached 56 in 1979, up from 27 in 1970.[6] The underlying causes for this growth process are very complex and have been analyzed in detail elsewhere*. Briefly, two factors appear to be crucial: the increased demand for higher education from the middle classes, and the effects of the economic crisis that has intermittently affected Mexico since the mid-1970s and has channeled a large number of young people who cannot find employment in the general labor market into the universities.

The economic crisis also produced a reduced rate of growth of medical services, which made it impossible for the health care system to absorb the increased number of graduates. As a consequence, new phenomena of un- and underemploy-

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Address reprint requests to Dr. Julio Frenk, Director, National Institute of Public Health, Apartado Postal 222, Admon. De Correos No. 1, 62000 Coernavaca Morelos, Mexico. Dr. Alagón is a Researcher with the Autonomous Technological Institute of Mexico; Nigenda is Head, and co-authors Muñoz-delRio, Robledo, Vaq́uez-Segovia, and Ramírez-Cuadra are researchers, all with the Department of Research on Human Resources for Health, Center for Public Health Research, National Institute of Public Health, Mexico. This paper, submitted to the Journal December 12, 1989, was revised and accepted for publication May 15, 1990.

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TABLE 1—Recent Evolution of Physician Supply in Mexico

Year	Absolute Number of Physicians	Physicians per 100,000 Population
1960	20,590	57
1970	34,107	71
1985	120,000	153

Sources:
For 1960: Ref. 7
For 1970: Ref. 8
For 1985: Ref. 9

ment began to characterize the medical labor market. Table 1 indicates that the supply of doctors in Mexico grew 2.7 times between 1960 and 1985.[7-9] It has been estimated that until 1971 every physician could find a salaried position upon graduation; by 1984 four graduates had to compete for every job in a health care organization.[10]

In addition, the number of positions for postgraduate residencies has remained practically constant throughout this period, at about 2,000 new positions per year. This means that the point of selection has been transferred from the medical school to the specialty residency, where entry is regulated by a uniform national examination, so that the competition for such positions has become fierce.

Since 1979 there has been a sustained effort to curb the growth of the medical education system, mostly through the institution of a joint decision-making body that brings together representatives from both the universities and the health sector.[11] As a result, only three new medical schools have been created in the past 10 years, and the annual number of graduates, which peaked at 14,099 in 1983, has declined steadily.[7] Nevertheless, the momentum of the physician population that was produced by the high growth rates of the 1970s explains why unemployment and underemployment have remained high.

Despite their visibility in the media, the magnitude of these imbalances has not been quantified with precision up to now. Estimates vary widely according to the interests represented by each source. In order to provide a scientifically validated picture of the employment situation of doctors, we conducted a national survey in 1986. Apart from its descriptive purpose, the survey also had the objective of exploring variables that may account for the performance of physicians in the labor market.

Methods

Data Source

Because the purpose of the study was to analyze the full range of employment possibilities, including the lack of a stable job, it was obvious that we could not draw a sample of traditional sites of medical work, such as offices and hospitals. Instead, we needed a representative sample of households where those holding a medical degree lived.

Such a sample was obtained from the 41,000 households that are visited every three months by the National Survey of Urban Employment, which is the official source of most labor statistics in Mexico and is carried out by the National Institute of Statistics, Geography, and Informatics. This survey covers the 12 metropolitan areas in the country and the four most important cities on the United States-Mexico border. It is a rotating panel survey, based on a stratified, multistage sample of households, where 88,050 people, representing 14,529,826 inhabitants ages 20 years or more, are interviewed. The validity and reliability of this survey have been extensively documented throughout the 17 years during which it has been in existence. Further details can be consulted in official publications on the survey.** Although this is a complex probability sample,[12] physicians constitute a relatively rare element in it; under these circumstances we assumed that they behave as a simple random sample, so that no weighting of the data is necessary in the analysis.[13]

In the last trimester of 1985, 620 physicians were identified during the regular round of the National Survey of Urban Employment. Shortly thereafter, these doctors were reinterviewed in face-to-face sessions by professional interviewers, who used a questionnaire specifically designed for our study at the National Institute of Public Health. The entire process of interviewing and coding was supervised by researchers from this institute.

Complete questionnaires were obtained from 604 physicians, for an overall response rate of 97 percent, which ranged from 94 percent in one city to 100 percent in eight cities. In Mexico future physicians are required, after four years of medical school, to complete a one-year internship in a hospital and one additional year of "social service" in a rural area before receiving their medical degrees. The sample included these interns and social service practitioners in addition to physicians who had already obtained their degrees. However, the present analysis is based on the 491 physicians who were active in the labor market. The "inactive" doctors excluded from the analysis comprise retired persons, as well as interns and social service practitioners, who do not demand employment since they are assured those positions in order to comply with the requirements for a medical degree.

Variables

Social Origin—This reflects not only the general standard of living that the physician enjoyed while he/she was growing up, but also the access to networks of relationships that may create future job opportunities. Social origin was measured through an index based on the occupation of the person who supported the physician during his/her years as a medical student (typically the father). The resulting categories were: low (e.g. manual workers, small-scale merchants), middle (e.g. technicians, teachers, clerical workers, foremen, sales agents, mid-sized merchants), and high (e.g. owners of industry, senior managers, senior government officials, professionals). In order to validate this index, we used the 1980 Census data to rank occupations along an income scale that was divided into three equal parts. The data on fathers' occupations from our survey were then matched to the Census occupation information. There was 100 percent match between our three categories of social origin and the three segments of the income scale.

Gender—Apart from its effects on productivity and job stability, gender is thought to be associated with differential access to occupational positions even af-

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**Instituto Nacional de Estadística, Geografía e Informática: Encuesta Nacional de Empleo Urbano: Diseño Muestral. Aguascalientes, México, December 1987.

ter educational opportunities have become more equal among women and men.

Medical School Quality—It cannot be assumed that attending a medical school represents the same educational experience for everyone. The only objective indicator of quality available for all medical schools in Mexico is a standard examination applied to graduates at the end of their studies in order to select those who will be admitted into a residency. However, this standard test has only been in existence since the mid-1970s, so that no comparable quantitative data are available for physicians who studied earlier. For this reason, we measured quality through an opinion poll from 30 experts in medical education selected to represent four groups: authorities from medical schools, decision makers from the main health institutions in charge of medical manpower, faculty members from medical schools, and academics who have studied various aspects of medical education. Experts were given a standardized schedule to rate the quality of schools, taking into consideration elements of structure (e.g. number and qualifications of the faculty, adequacy of the facilities), process (e.g. type of curriculum, educational methodology), and outcomes (e.g. competence of graduates, success in finding jobs). They were asked to provide separate evaluations for two time periods: before and after 1970, since it cannot be assumed that the quality of the schools existing before 1970 has remained constant. Evaluations for each school in each time period were made on a scale from 1 to 4, and the scores were weighted according to the amount of information that the experts declared themselves to have about each assessment. The quality index was the average of these weighted scores. Its distribution revealed two marked cleavages, which were used as cutoff points to stratify schools into three categories of quality: inadequate, average, and good. The validity of the quality index for schools after 1971 was checked by correlating it with the average scores of graduates from each school on the standard examination for admission into residencies. The Pearson r correlation was 0.76, which was considered satisfactory.

Type of Specialty—This variable reflects the differential access to opportunities for postgraduate education. It also indicates the intensity of specialization, as measured through the following categories: no specialty, family medicine, basic specialties (i.e. internal medicine, pediatrics, gynecology-obstetrics, and general surgery), and subspecialties. Even though in Mexico the length of postgraduate training is similar for family medicine and for the basic specialties, the former is considered to represent a lower degree of specialization, since family practitioners are trained to deal with all age and sex groups. In addition, the practice settings are also different. Family medicine emerged during the 1970s as a residency program aimed at satisfying the requirements for first-contact care at the social security and other public-sector organizations. Thus, family doctors practice almost exclusively in ambulatory settings, while basic specialists also work in general hospitals.

Generation—To consider the influence of contextual variables, such as the situation of the medical labor market, we measured a variable called generation, which refers to the year in which the physician entered medical school, whether before or after 1971, the year when the policy of massive access to higher education was fully implemented. In addition to reflecting the environment to which the physician was exposed during his/her formative years, generation is very highly correlated (Pearson's r = 0.97) with age. In turn, age is related to experience, productivity, and exposure to employment opportunities, all of which affect the success of a physician in obtaining a desirable job.

Even though the data come from a cross-sectional survey, the fact that there is an unambiguous chronological order among the variables allows us to adopt a "biographical" view of medical employment. In this time order, three of our independent variables can be considered antecedent to the others, i.e. social origin, gender, and generation/age. In turn, attending a medical school occurs before postgraduate training to become a specialist. The final link in this chain, which represents the dependent variable, is medical employment.

Dependent Variable—We conceptualize the dependent variable as "medical employment pattern," built through cluster analysis of physicians sharing similar job characteristics. The detailed statistical procedure has been described elsewhere.[14] The final variable summarizes the most relevant dimensions of employment: number of jobs; position (salaried or independent); occupational level (whether the job requires less skill than what the person was trained for, e.g. whether the physician performs a nonmedical job); productivity (measured by the number of patients seen and hours worked per week); income; and job stability.

Obviously, the number of potential combinations is very large. Cluster analysis enabled us to ascertain the specific combinations that occur empirically. These combinations were grouped to form the dominant patterns, which turned out to be six:

1. Unemployment: This category was strictly defined to include those doctors who did nothing but search for a job or who had stopped looking for one because they had become discouraged.

2. Qualitative Underemployment: This pattern refers to physicians who worked outside medicine. Even if they may have had high productivity or income, the social investment in their medical education would have been underutilized. As it turned out, most of these jobs required a lower level of training than medical care.

3. Quantitative Underemployment: Our analysis identified a cluster formed by physicians who provided patient care, but did so under one or more negative circumstances suggesting that their work potential was severely underutilized.[15] As shown in Table 2, doctors in this cluster saw very few patients (the mean for the cluster was 18.2 patients per week, compared to 66.1 among all other physicians in patient care), worked reduced hours (a mean of 31.9 hours, compared to 55.3 among the rest of physicians, with 60 percent of underemployed doctors working less than 35 hours per week), and/or earned a low income (54 percent made less than three minimum wages, compared to 18 percent among fully employed doctors). In addition, these physicians tended to have unstable working conditions, characterized by job insecurity and high turnover rates.

4. Salaried Single Medical Employment: This is a conventional pattern composed of doctors with one stable salaried job in a health care organization, most often in the public sector.

5. Independent Single Medical Employment: This cluster represents another conventional pattern, wherein physicians work full time in their own private practices. In terms of income, this was a more heterogeneous cluster than the previous one; as shown in Table 2, the percentages in the lower end of the distribution were similar among both clusters with single employment, but independent doctors exhibited a much higher proportion of high-income earners. Such heterogeneity also accounts for the smaller average produc-

TABLE 2—Average Productivity and Income for Medical Employment Patterns Involving Patient Care

Medical Employment Patterns Involving Patient Care	Productivity: Mean number of patients seen per week	Productivity: Mean number of hours worked per week	Income: Percentage earning less than 3 minimum wages	Income: Percentage earning more than 5 minimum wages
			%	%
Salaried single employment	65.4	60.1	25	35
Independent single employment	27.9	36.7	24	53
Multiple employment	83.8	57.1	5	84
Quantitative underemployment	18.2	31.9	54	19

TABLE 3—Relationships between Medical Employment Pattern and Each of the Independent Variables

	Independent Variables															
	Generation		Gender		Social Origin			Medical School Quality			Type of Specialty				Total	
Medical Employment Pattern	1920–70 %	1971–82 %	Male %	Female %	Low %	Middle %	High %	Inadequate %	Average %	Good %	Without Specialty %	Family Medicine %	Basic Specialties %	Subspecialties %	%	N
Salaried single employment	27	41	32	45	37	38	25	33	38	32	25	83	37	39	34	169
Independent single employment	16	7	13	4	11	11	12	11	11	11	15	0	5	11	11	55
Multiple employment	42	10	30	9	20	27	33	9	27	40	14	15	48	41	26	126
Quantitative underemployment	6	15	9	15	10	11	12	18	9	6	15	0	8	5	11	52
Qualitative underemployment	7	14	11	10	11	9	11	14	10	9	18	2	1	3	11	52
Unemployment	2	13	5	17	11	4	7	15	5	2	13	0	1	1	7	37
Total %	100	100	100	100	100	100	100	100	100	100	100	100	100	100	100	
N	235	256	391	100	197	180	114	158	166	167	262	41	89	99		491
X^2	98.5		40.8		16.9			65.4			159.4					
p	0.0001		0.0001		0.0759			0.0001			0.0001					
Cramer's V	0.45		0.29		0.13			0.26			0.32					

tivity, since competition probably forces some independent physicians into the lower end of the distribution.

6. *Multiple Medical Employment:* This last pattern included cases holding two or more medical positions, the majority of them with one salaried and one independent job. In the Mexican medical labor market, as in many other developing countries, such a dual pattern has been traditionally considered to be very desirable, since it allows the physician to combine the prestige of a position in a public hospital with the economic incentives of a private practice. The advantages of this situation can be appreciated in the productivity and income figures presented in Table 2.

Analysis

The purposes of the analysis were to measure the frequency of each medical employment pattern, to determine the links among the variables, and to compare the relative strength of each relationship. In order to achieve this last purpose, we used percentage differences, as well as two conventional measures of association for cross-tabulations, namely, Cramer's V when the measurement level of the variables was nominal and Gamma when it was ordinal.[16–18]

Results

Table 3 presents a set of cross-tabulations between pairs of variables. The marginal totals provide the frequency distribution of employment patterns, and the percentages can be interpreted as employment rates. Thus, the unemployment rate is 7 percent, while the global underemployment rate (quantitative and qualitative) is 22 percent. Projecting these proportions, we can infer from the population of active physicians in the 16 cities (81,024) that un- and underemployment affect approximately 23,500 doctors.

Moving to the correlates of this problem, Table 3 indicates that all the explanatory variables, except social origin, exhibit significant and fairly strong overall associations with employment pattern. Beyond these common characteristics, it is worthwhile to examine each cross-tabulation.

Generation

Table 3 shows a marked increase in unemployment and underemployment from the oldest to the youngest group. Among those who do have a job, there is a shift from independent to salaried practice and from multiple to single employment. This variable reflects not only the intrinsic differences between the two generations, but also the different time spans that each group has been in the labor market. In addition, the sample may be biased for the older generation insofar as we only have the survivors. Nevertheless, the trends exhibited by our data are consistent with existing information about the supply and demand of physician manpower.

Gender

There is an important connection between generation and gender, since the proportion of women among physicians was 14 percent in the 1920–70 generation and 26 percent in 1971–86. As can be seen in Table 3, the proportion of unemployed women is three times higher than the corresponding figure for men. Although salaried single employment is the most important pattern for both groups, the second most important pattern is multiple employment among men and unemployment among women.

Social Origin

There is a tendency for multiple medical employment to be more frequent with higher social origin. Conversely, salaried single employment is not as important in the high group as it is in the other two, and unemployment is most frequent among physicians from the lowest socioeconomic background.

Medical School Quality

There is a marked descending gradient for the least desirable patterns as quality increases: 47 percent of physicians who studied at an inadequate school are un- or underemployed, compared to 24 percent and 17 percent of graduates from average and good schools, respectively. It is also worth noting that higher quality is associated with increasing proportions of multiple medical employment. This is the dominant pattern for physicians from the better schools, whereas salaried single employment is the most common pattern for inadequate and average schools.

Type of Specialty

Lack of postgraduate training greatly increases the chances of un- or underemployment: almost half of the physicians without a specialty are located in one of these categories. The practice of family medicine has become the prototype of medical work based on a single salary at the service of the state, with 83 percent of family doctors exhibiting such a pattern. The traditional pattern of multiple employment is now reserved for physicians with a basic specialty or a subspecialty, among whom unemployment is very rare.

Some of the effects shown in Table 3 could be indirect, as mediated by chronologically subsequent variables. In the case of gender, there were no significant differences between men and women in the quality of the medical school attended, but their respective opportunities for postgraduate training presented sharp contrasts. Thus, when specialty is examined by sex, we find that more than two-thirds of female doctors have no specialty, compared to less than half of their male colleagues (Table 4). Among those who do complete a residency, the proportions of men and women are about equal in family medicine. However, males dominate in the basic specialties and even more so in the subspecialties, where their proportion is twice that of females.

Social origin could also affect employment indirectly, through a relationship that is mediated by medical school. As shown in Table 5, the association between social origin and medical school is fairly strong. Thus, 40 percent of physicians from the lower social origin attended an inadequate medical school, compared with 44 percent of doctors from the higher social origin who went to a good school.

Part of the effect of medical school quality on employment may be mediated by type of speciality, since entrance to a residency is a competitive process that is bound to be affected by the quality of previous education. Table 6 demonstrates that there is a strong association between medical school quality and type of specialty. The likelihood of having no specialty increases consistently as quality decreases. Thus, the proportion of general practitioners is 1.4 and 2.2 times larger among graduates of inadequate schools than among doctors who studied at average or good schools, respectively. In contrast, good schools concentrate the highest proportions of basic specialists and subspecialists. In fact, we can consider the high degree of consistency of this association as a confirmation of the validity of our measure of school quality.

Discussion

The results of this study point to a series of imbalances, inadequacies, and inequities that characterize the medical labor market in Mexico. First of all, the heavy burden on Mexican society: almost one-third of all physicians living in the 16 major cities are unemployed or underemployed. Because these extreme forms of imbalance are likely to be concentrated in the large cities, the situation probably approaches the total for the whole country. In any event, the figure of 23,500 unemployed and underemployed physicians acquires its precise meaning when one considers that in 1986, the year when our survey was conducted, 9.3 million people, representing 11 percent of the Mexican population, did not have ready access to health services.[19]

TABLE 4—Type of Specialty by Gender

Type of Specialty	Gender Female %	 Male %
Without specialty	69	49
Family medicine	9	8
Basic specialties	11	20
Subspecialties	11	23
Total	100	100
N	100	391

$X^2 = 14.56$ $p = 0.0022$ Gamma = 0.36

Not all types of imbalance produce the same social impact. Unemployment and qualitative underemployment represent a waste of resources for society and a source of frustration for those doctors who have to work outside medicine or cannot find a job at all. Paradoxically, even those who are unemployed seem unwilling to serve in deprived rural areas, indicating that they would rather remain in the cities than be fully employed as physicians where their services are most needed.[20]

While unemployment and qualitative underemployment are undoubtedly severe problems, quantitative underemployment opens up a new perspective on physician oversupply. In this case, doctors still provide patient care, but do so at low levels of productivity. It is not unreasonable to assume that such doctors may represent a potential risk to their patients. Our analysis indicates that physicians with

TABLE [illegible]—Quality of the Medical School Attended, by Social Origin of the Physician

Medical School Quality	Social Origin Low %	 Middle %	 High %
Inadequate	40	28	25
Average	36	34	31
Good	24	38	44
Total	100	100	100
N	197	180	114

$X^2 = 17.22$ $p = 0.0018$ Gamma = 0.24

TABLE 6—Type of Specialty, by Quality of the Medical School Attended

	Medical School Quality		
Type of Specialty	Inadequate %	Average %	Good %
Without specialty	75	52	34
Family medicine	6	9	9
Basic specialties	10	19	26
Subspecialties	9	20	31
Total	100	100	100
N	158	166	167

$X^2 = 58.97$ $p = 0.0001$ Gamma = 0.45

low output tend to be young, have no postgraduate training, and often work in independent practices without any supervision. Under these circumstances, the quality of care may suffer through two mechanisms: first, doctors who see few patients may be unable to build up enough clinical experience; second, there may be a tendency to overprescribe medical procedures in order to derive a living from those few patients. The extent to which quantitative underemployment of doctors may represent a health risk is a subject for future research.

The appearance of imbalances also suggests that the medical labor market is experiencing a process of segmentation.[21] Apart from those physicians who are out of a medical job, there seem to be two major segments. The first one is characterized by reasonable productivity, relatively sufficient earnings, and stable employment conditions through a salaried position in government, a conventional private practice, or both. Diverging from the mainstream, a substantial number of doctors have entered into a marginal segment of the market, which is characterized by low productivity, insufficient earnings, and occupational instability. The segmentation process even extends to those who have attained relatively good positions, as illustrated by the fact that the traditional prestigious pattern of combining a salaried position in the public sector with a private practice is now dominated by physicians with a basic specialty or a subspecialty; in contrast, family doctors exemplify a new kind of medical functionary, since the vast majority of them hold their only job at the service of the government.

These cleavages in the medical profession reflect a broader process leading the medical education and health care systems to reproduce, rather than ameliorate, pre-existing social inequalities. Such a process begins with the channeling of students from different social origins to medical schools of varying quality. Indeed, the data reported here have demonstrated that there is a substantial degree of social stratification among medical schools, thus corroborating the findings from a previous study.[22] In turn, the quality of the medical school affects the probability of doing a specialty residency, which has a direct effect on employment. To complete this logic, we can assume that underemployed physicians are more likely to care for people from the lower social classes. There are reasons to believe that the quality of the care provided by these physicians is inadequate. In addition, there is empirical evidence that the distribution of the quality of care in Mexico is highly stratified by social class.[23] A circle is therefore closed: disadvantaged people will attend substandard medical schools and end up providing inferior care for other disadvantaged people. Obviously, this is not a deterministic but a probabilistic sequence of events. Indeed, we are not dealing with a rigid social structure that prescribes these patterns of behavior, but with one that tends to reproduce inequality. Although many individuals can and do escape from this vicious cycle, its existence seriously threatens the legitimacy of higher education and of health care.

The distribution of employment patterns by gender also exhibits important elements of inequality. In contrast to social origin, there is no stratification of medical schools by gender. Yet disparities are found in the opportunities for postgraduate residency, as shown by our data. Most of the substantial gender differentials in employment patterns can be attributed to the mediating effect of access to residencies. Such differentials acquire additional importance when one considers that the proportion of women in the medical profession is increasing. For example, by 1985 women constituted half the entering class at the most important medical school in the country, up from 38 percent in 1979.[24]

Despite their many sobering implications, our findings still leave ample room for policy interventions to correct imbalances. Many of the effects on employment appear to be mediated by the quality of medical schools and the opportunities for postgraduate training. Thus, actions can be implemented to improve inadequate schools and to make admissions dependent on academic potential only and not on the economic status of the student's family. Similarly, residency programs can adopt affirmative action measures that give women equal opportunities based solely on qualifications. In particular, it seems necessary to provide support services that will make the pursuit of a medical career compatible with other social roles.

With respect to those physicians who are already un- or underemployed, it would seem urgent to adopt continuing education programs that may compensate deficiencies in previous training and encourage those physicians to serve in rural areas. Such educational endeavors would have to be complemented by specific incentives, both in monetary terms and in opportunities for future career advancement. The ultimate obligation of protecting the health of the population also makes it necessary to consider mechanisms for the periodic certification of medical competence.

Finally, current effects to restrict the growth of medical schools must continue. Indeed, there can be little doubt that an overabundance of physicians has dysfunctional consequences not only for the medical profession itself, but also for the health system as a whole, including the users of services.

Above all, countries like Mexico must make explicit efforts to redress the inequities that still mar the systems of higher education and health care. The French sociologist Raymond Boudon spoke of "perverse effects" as an analytic category that accounts for the actual functioning of many social institutions.[25] Higher education and health care, two of the fundamental vehicles for promoting equity, are at risk of falling into the perverse effect of reproducing social inequalities. Unless they truly fulfill their role of ameliorating the negative consequences of poverty, they may themselves end up being impoverished institutions. The challenge, in essence, is to harmonize the legitimate aspirations for social advancement through higher education with the imperative of improving the well-being of the population. ☐

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Why Do Child Mortality Rates Fall? An Analysis of the Nicaraguan Experience

ABSTRACT

A comprehensive review of available sources of mortality data was undertaken to document the changes that have occurred in infant mortality in Nicaragua over the last three decades. It was found that a rapid fall in infant mortality commenced in the early 1970s and has continued steadily since. Trends in several different factors which might have led to this breakthrough were examined including: income, nutrition, breastfeeding practices, maternal education, immunizations, access to health services, provision of water supplies and sanitation, and anti-malarial programs. Of these, improved access to health services appears to have been the most important factor. At a time when the number of hospital beds per capita was dropping, increasing numbers of health care professionals, particularly nurses, were becoming available to staff primary health care facilities built in the 1960s. These were provided at least partly in response to the growing political turmoil enveloping the nation at that time. Certain Nicaraguan cultural attributes may have added to the impact of the reforms. Efforts in the field of public health made since the 1979 insurrection appear to have maintained the decline in child mortality. (*Am J Public Health* 1991; 81:30–37)

Peter Sandiford, MBChB, MMedSci, Patricia Morales, BA, Anna Gorter, MD, Edward Coyle, MBChB, and George Davey Smith, MB BChir, MA, MSc

Introduction

Poor child health in developing countries continues to be a source of concern.[1] In recent years it has become apparent that while poverty is probably the major factor affecting survival in the early years of life, there are undoubtedly other influences playing an important role. Some countries, in spite of a low per capita gross national product, have succeeded in achieving mortality levels close or equal to those of the industrialized world. In contrast, there are others with poor levels of child health despite enjoying relatively high per capita incomes.[2] Efforts have recently been made to identify the factors beside socioeconomic status which are determinants of infant mortality. In 1985, the Rockefeller Foundation sponsored a conference which studied the striking mortality reductions in Kerala State (India), Sri Lanka, Costa Rica, and China.[3] In each case, mortality has fallen sharply over a relatively short period leading to high life expectancy relative to national income. Seven conditions were purported to be essential to the observed improvements: female autonomy, education (especially female), the provision of accessible health services, a mechanism to guarantee the efficient operation of health services, an adequate minimal standard of nutrition, universal immunization, and a commitment to the establishment of effective antenatal and obstetric services.[2]

Nicaragua is another country which has experienced a recent breakthrough in child mortality. This article describes that breakthrough and looks at its relationship with some of the factors alleged to be important determinants of child mortality. It is hoped this work serve as a model for the type of intersectoral data analysis advocated recently by de Kadt.[4]

Methods

Seven different sources of data were identified which could provide time-located estimates of infant mortality in Nicaragua from 1960 to 1986.[5–11] A brief description of each source of data is given in the Appendix. Except for data obtained from the Civil Register, the infant mortality estimates were derived using the Brass technique based on the ratio of surviving to ever born children of mothers classified by five-year age group. The rationale behind this method is that the proportions dead of ever born children for women of a certain age correspond to discreet child survival probabilities. These survival probabilities can be used to calculate infant mortality rates (IMRs) if one assumes a typical age pattern of mortality. In this case the Coale-Demeny 'West' model of mortality was used.[12]

Variations of the Brass method have been developed which allow the estimation of trends in mortality change, recognizing that the deaths in children born to

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Address reprint requests to Dr. Peter Sandiford, Department of International Health, Liverpool School of Tropical Medicine, Pembroke Place, Liverpool L3 5QA, England. Ms. Morales is with the Programa de Ingenería Ambiental, Universidad Nacional de Ingenería, Managua, Nicaragua; Dr. Gorter is with the Ministry of Health, Managua; Dr. Coyle is with South Glamorgan Health Authority, Cardiff, Wales; Dr. Davey Smith is with the London School of Hygiene and Tropical Medicine, England. This paper, submitted to the Journal January 19, 1990, was revised and accepted for publication August 13, 1990.

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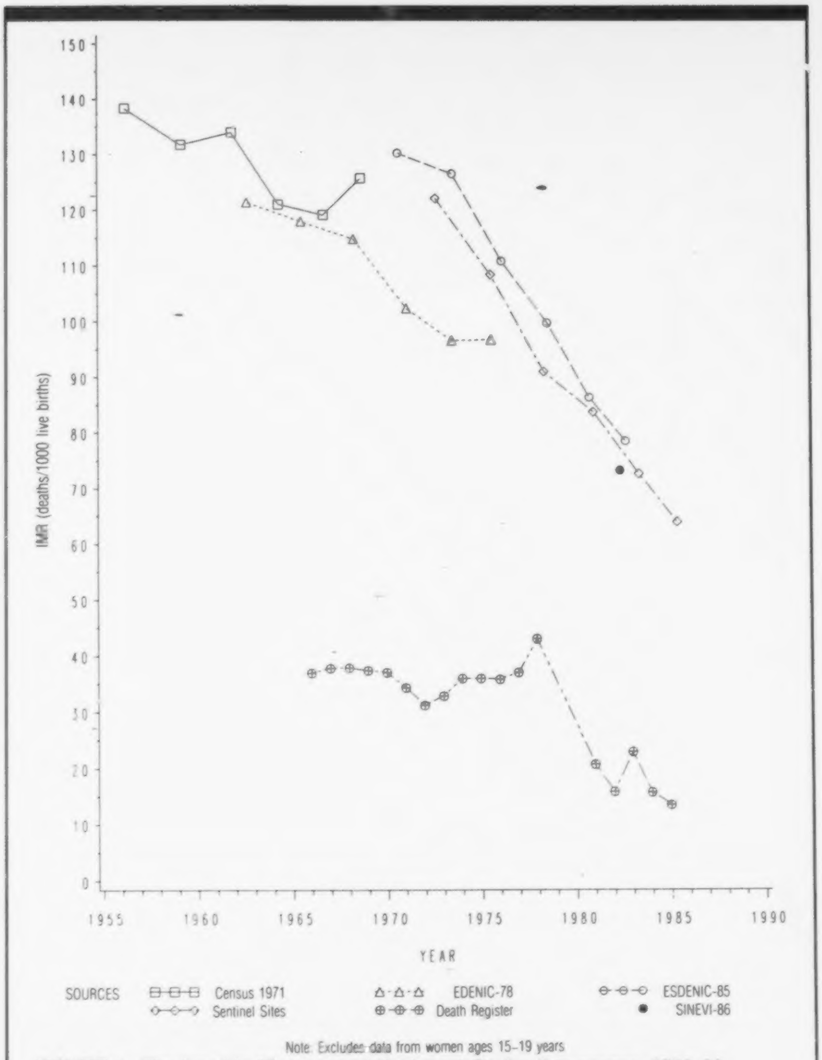


Note: Excludes data from women ages 15–19 years

FIGURE 1—The trend in Nicaraguan infant mortality by data source, 1956–86

women ages 15–19 years have occurred more recently than the deaths in children of women ages 45–49. The Trussell method, which was employed in this analysis is one such variant.[13] It provides time locations for the infant mortality estimates obtained from each five-year age group of mothers.

Estimates of infant mortality from the death register were calculated by dividing the number of deaths registered in each year by the projected population under the age of one. Death registration is well known to be variable and incomplete so these figures were not used in further calculations. However, the relatively low rate of omission for birth registration allows a more accurate estimation of the IMR by asking about the survival of preceding births at the time of registration. The data for 1986 (SINEVI-86) were analyzed in this way by the Centro Latino-Americano de Demografía (CELADE).[11]

There are a number of factors which are believed to be important determinants of child mortality. These include: income, nutrition, maternal education, breastfeeding, immunization coverage, the provision of health services, access to water supplies and sanitation, malaria control, as well as certain cultural attributes of the population. Trends in these factors were examined and related to the observed changes in infant mortality in Nicaragua from 1960 to 1989. Wherever possible, data were obtained from the original sources in Nicaragua by visiting libraries and documentation centers in government ministries and institutions.

Data on per capita Gross National Product (GNP) were obtained from an analysis performed by the National Autonomous University of Nicaragua.[14] It was not possible to accurately quantify changes in income distribution.

An assessment of trends in child nutritional status was based on three surveys on nationwide samples of children in 1966, 1977–8, and 1986.[15–17] Height for age was used as the indicator of nutritional status as it was the only one measured in all three surveys. In two of the surveys,[15,16] the children measured were under the age of 5, in the third they were school children ages 6–9 years. Figures for food production and caloric intake were derived from World Bank[18] and United Nations Food and Agricultural Organization[19] statistics, respectively. The frequency of breastfeeding in 1977 and 1988 was estimated from two ad hoc surveys performed in Managua.[20,21]

Levels of illiteracy in Nicaragua were obtained from the censuses of 1950,[22] 1963,[23] and 1971[5] and from the EDENIC-78 and ESDENIC-85 surveys.[11] Rates were calculated for the total population ages 10 years and over, for the urban and rural populations, and for women ages 15–44 years.

An analysis of Ministry of Health cause of death statistics by CELADE documents changes in death rates in children due to immuno-preventable diseases.[11] These were compared with an indicator of vaccination coverage, estimated by dividing the Ministry of Health figures for the annual number of inoculations by the population of children under the age of five for that year.

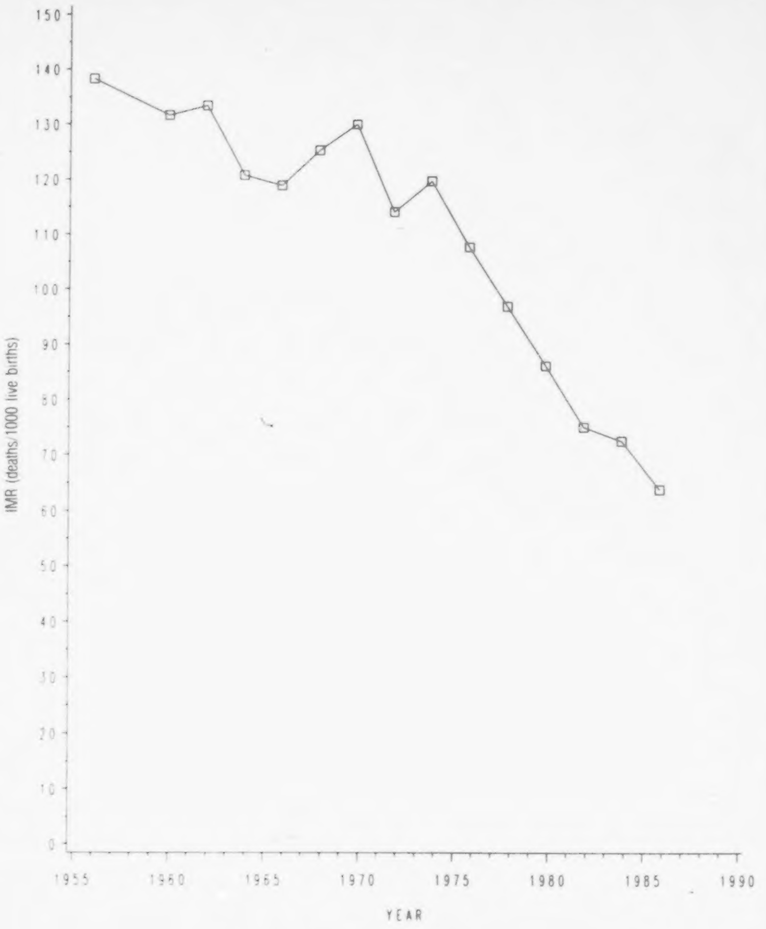
There have been several analyses of Nicaragua's health system, mostly comparing the situation before and after the 1979 insurrection.[24–27] These were reviewed and combined with statistics from the health sector[10,28–30] (public and private) to obtain an overview of the trends in the provision of health services—both as infrastructure and as human resources—from 1960 to 1985.

National statistics were used to determine what changes had taken place in domestic water supply and sanitation facilities.[10,30] Figures for latrine construction were taken from Ministry of Health records.[28] An analysis by Garfield and Vermund[31] and national records of laboratory tests for malaria were used to look at trends in the incidence of malaria.

Results

Infant Mortality Estimates

Figure 1 shows the infant mortality estimates from each of the aforemen-



Sources: Census 1971, EDENIC-78, ESDENIC-85, SINEVI-86, Sentinel Sites 1988

FIGURE 2—The trend in Nicaraguan infant mortality: 1956–86 (weighted averages)

tioned sources plotted against time. Estimates from ESDENIC-85, SINEVI-86 and the 1971 Census, the three largest databases, give a consistent profile of the trend in mortality. Moreover, the IMRs from the sentinel sites study agree remarkably well with those of ESDENIC-85. There is some divergence in the mortality estimates from EDENIC-78, particularly for the early 1970s but these are based on data from a small number of women and hence this may be due to sampling error. The rates calculated from death registration are incompatible with the other estimates, a reflection of gross inadequacies in coverage of this system. They are not used in further analyses.

A weighted average of infant mortality estimates from the 1971 census, the EDENIC-78 and ESDENIC-85 surveys, the SINEVI-86 analysis of birth registration and the sentinel sites study was made for each two-year interval. Rates obtained from women ages 15–19 years were not included as these are well known to be unreliable.[13] The results are plotted in Figure 2. Figure 3 presents the same data plotted on a logarithmic scale which more appropriately portrays the change in the rate of decline. Taken as a whole, the studies indicate that the rate of fall in infant mortality, which was slow in the 1960s, increased sharply in the early 1970s, and that this rapid decline was still continuing in 1986. The infant mortality rate drops on average by 4.7 deaths per 1,000 live births for each year from 1974 to 1986. Assuming a 'West' mortality schedule, this is equivalent to an annual increase in life expectancy of 0.75 years each year over this 12-year period.

Figure 4 presents urban and rural IMRs from the 1971 Census and ESDENIC-85. Changes in the definitions of 'urban' and 'rural' and internal migration make interpretation difficult but there is no clear evidence that in Nicaragua the breakthrough was more pronounced in rural areas than in urban areas.

Differential infant mortality rates according to levels of maternal education are provided by the CELADE analysis. In the three time periods examined—1966–67, 1973–74, and 1982–83—infant mortality rates for mothers with secondary or higher education were considerably lower than for mothers with primary or no formal education. The results suggested that this differential narrowed over the period in question (data not shown).

Explanations for the Decline in Child Mortality

How can this sharp drop in child mortality be explained? In the first place, it should be recognized that more than one factor may be involved. The reason that mortality started to fall may be different from the reason that it continues to fall. Since this is not a controlled experiment, causes can easily be confused with effects, and one is always left in doubt as to what would have happened in a different set of circumstances. Nevertheless, the Nicaraguan case has some very interesting features which make it worth examining in greater detail.

Income—The observed infant mortality decline in Nicaragua does not correspond to a rise in general living standards. Per capita GNP rose rapidly in the early 1960s but flattened off in the 1970s before falling sharply in the 1980s. This fall may have been offset somewhat by a more equal distribution of wealth following the insurrection (as a result of the wages policy, food subsidies and rationing, price controls, and rent controls), but there are insufficient data to quantify changes in income distribution over this period.

Nutrition—In 1966, 36 percent of children under the age of 5 were found to be below two standard deviations in the reference range for height for age. In the survey performed from 1977 to 1978, 26 percent of children under the age of 5 were below two standard deviations for height for age and in 1986 the corresponding prevalence of stunting in first grade school children (ages 6–9 years) was 22 percent. Although Nicaragua has high rates of primary school attendance, poor children tend to start later than the rest. This produced an age bias in the results of the 1986 survey such that in 6-year-old first grade children the prevalence of low height for

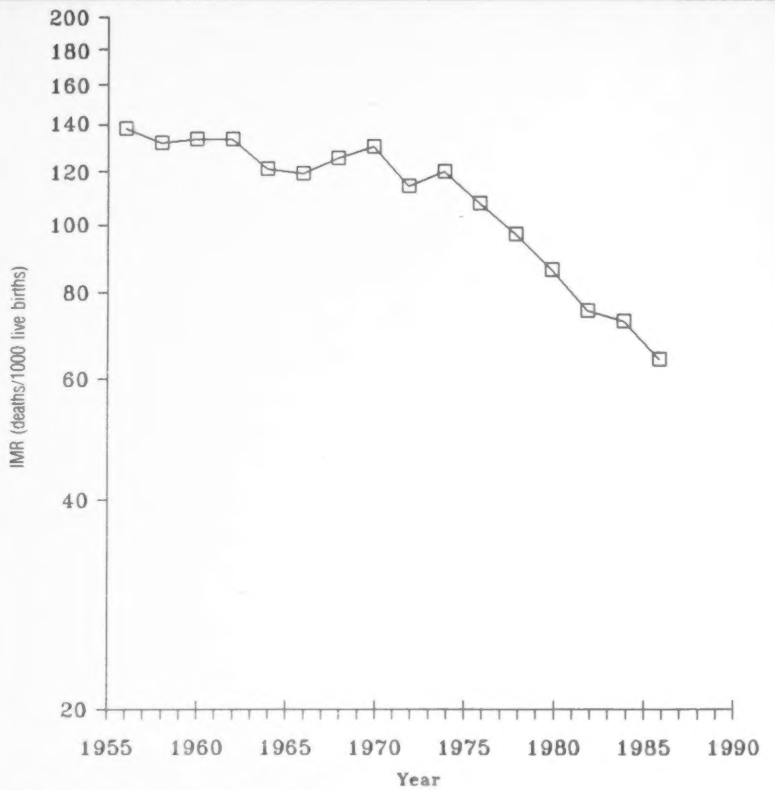


FIGURE 3—The trend in Nicaraguan infant mortality: 1956–86 (weighted averages) plotted on a logarithmic scale (base 10)

age was only 14.5 percent compared with 32.7 percent for 9-year-old first grade pupils.

Despite the deficiencies in the available data, it would seem reasonable to conclude that nutritional status in Nicaraguan children improved steadily from 1966 to 1986, but the factors leading to this improvement are uncertain. Food production in 1986 was only 70 percent of what it was in 1967 having fallen sharply from its peak in 1978. The number of food calories per head of population dropped from 2,410 in 1970 to 2,344 in 1975 and then to 2,258 in 1980, recovering slightly by 1983 to 2,304.

On the other hand, a variety of programs and activities aimed at improving child nutrition already existed in 1970.[32] These were expanded in 1973 in response to the threat posed by the 1971 drought and the 1972 earthquake. A milk distribution program of the Ministry of Public Health was in operation since the 1960s but was gradually abandoned in the late 1970s. It was recommended for several years following the 1979 insurrection along with other policies aimed at establishing an adequate nutritional standard. These included food subsidies, rationing, and supplementary feeding programs for pregnant women and malnourished children.

Breastfeeding—The limited data available suggest that there has been no increase, and perhaps a decrease occurred in the average duration of breastfeeding in Nicaragua over the period of rapidly declining child mortality. In 1977, 58 percent of mothers in Managua were breastfeeding their children for at least six months. In 1988, only 33 percent of children between three and five months of age were still receiving breast milk.

Maternal Education—The changes in the rate of illiteracy in Nicaragua are summarized in Table 1. The tendency has been for illiteracy to decrease throughout the period from 1950 to 1985. The overall drop in illiteracy is greater than the individual falls in urban and rural areas because of migration to the town and cities. The slowest rate of change in the overall prevalence of illiteracy was observed between 1971 and 1978. Between 1978 and 1985 there was a marked drop in the prevalence of illiteracy, mainly as a consequence of the 1980 National Literacy Crusade and the extensive adult education program which followed.

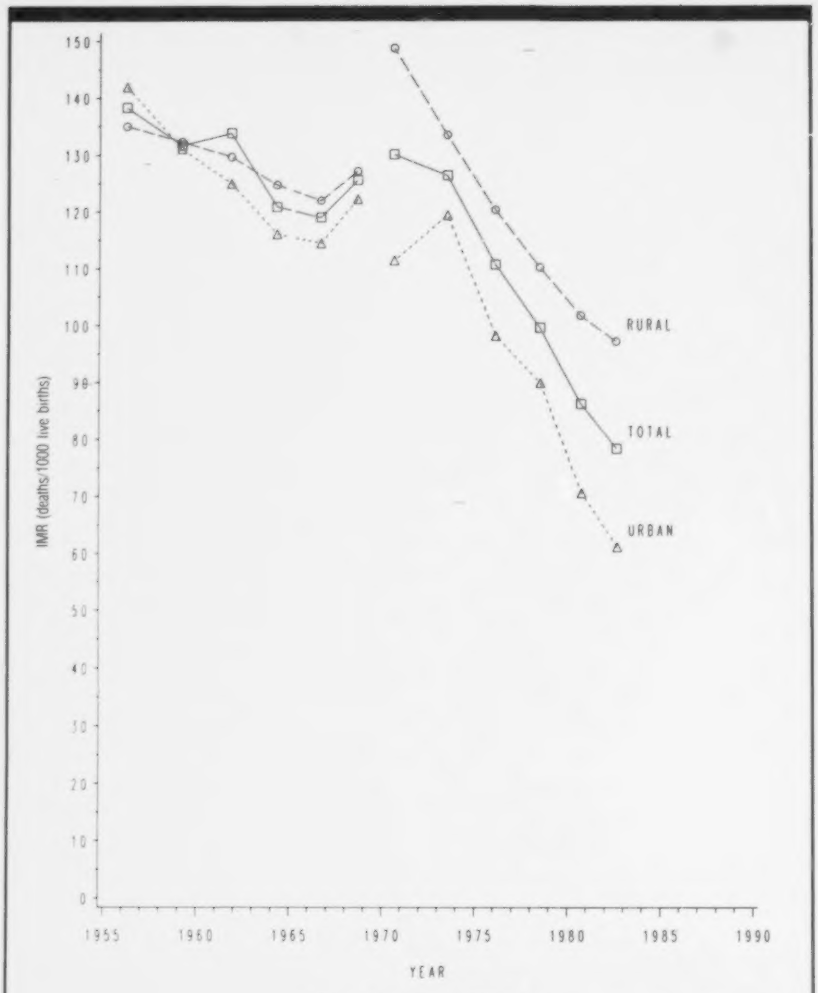
Immunizations—Mortality due to diseases preventable by immunization in the ages 1–4 dropped from 21 per 10,000 in 1975 to 3 per 10,000 in 1986 (or from 17 percent to 4 percent as a proportion of total mortality in this age group). Unfortunately, the poor quality of the statistics did not permit a similar analysis for children under one year of age.

Table 2 shows the change in rates of immunization over the period from 1970 to 1985. Prior to 1979 rates of vaccination coverage varied widely as periodic campaigns were launched in response to cyclic epidemics of immuno-preventable disease. Since 1979 coverage has been higher and shown less variation. The participation in massive nationwide vaccination campaigns of large numbers of voluntary health workers in the years immediately following the insurrection was probably the main factor in this improvement.

Health Services—In comparison with other developing countries, Nicaragua's health system suffered more from maldistribution of resources than from a lack of investment. Health service infrastructure and personnel were concentrated in the cities and particularly in the capital. In the 1960s and 1970s, efforts were made to redress this imbalance and improve access to health care. A USAID loan of $US 2.2 million was received in 1965 to set up 55 health centers and provide outreach programs in rural areas. By 1974 there were 33 hospitals, 93 health centers and 38 health posts spread throughout the country, and the Rural Mobile Health Program provided care to remote areas. More primary health care facilities were built in the first years following the insurrection.

Although considerable infrastructure was already installed by 1973, an evaluation of the 55 health centers built from USAID loan found low utilization rates owing to the lack of personnel, supplies, and management support. Two further USAID grants were made in the mid-1970s to correct these deficiencies.

In contrast to the growth in primary health care facilities, there was a general deterioration in hospital care since the early 1970s. The number of hospital beds per capita peaked at 2.6 in 1971 and has gradually dropped since. The National Social Security Institute (INSS), a health care provider, grew slowly but steadily from 1960 to 1980. In 1963, only 4.1 per-



Sources: Census 1971 and ESDENIC-85

FIGURE 4—The trend in Nicaraguan infant mortality: Urban versus rural, 1956–82

TABLE 1—Trends in Nicaraguan Illiteracy Rates, 1950–85

Year	Total[a]	Urban[a]	Rural[a]	Women[b]
1950	62%	30%	80%	61%
1963	49%	21%	70%	50%
1971	42%	18%	66%	43%
1978	39%	17%	63%	38%
1985	25%	13%	40%	26%

[a] Refers to the population ages 10 years and over.

[b] Refers to the population ages 15–44 years.

Sources: 1950 Census,[28] 1963 Census,[29] 1971 Census,[6] EDENIC-78,[12] ESDENIC-85 (unpublished data).

cent of the population were insured. By 1971 coverage of the population had increased to 7.5 percent and by 1978 to 8.4 percent. In 1980 under the Government of National Reconstruction, coverage reached 10.1 percent of the population.

The number of doctors and especially nurses per head of population increased between 1974 to 1985. From 1970–74, there were 5.4 doctors and 2.3 nurses per 10,000 inhabitants. For the period 1975–79 there were 5.8 doctors and 2.6 nurses per 10,000 population and in the five years from 1980–84 there was an average of 6.0 doctors and 3.3 nurses per 10,000 population. The impact of the increased availability of health care professionals (together with dropping numbers of hospital beds per head of population) was reflected in rising per capita rates of hospital discharge and rapidly increasing numbers of outpatient consultations (especially emergency consultations).

Other Public Health Measures—Nicaragua undoubtedly had low rates of access to plentiful clean water supplies and sanitation prior to 1974, but the situation has changed little since then. In 1981, the number of urban domiciliary connections was almost identical to that prior to the earthquake in 1972. Latrine construction by the Ministry of Health, which averaged 4,590 per year from 1974–78, was barely sufficient to keep up with the increase in the rural population.

The use of residual insecticides for the control of malaria decreased steadily in the late 1970s as the World Health Organization campaign for the elimination of malaria lost its momentum. The number of notified cases varied widely from one year to the next but showed no secular downward trend from 1972–82. The average annual reported number of laboratory confirmed cases was 15,386 from 1972–76 and 18,738 from 1977–81. A renewed effort was made to control malaria in 1981 by the administration of mass chemoprophylaxis but the impact was transient.

Discussion

From a variety of sources of infant mortality estimates, this paper has documented a rapid decline in infant mortality in Nicaragua which commenced in about 1974 and was still in progress in 1986. The rate of decline since 1974 is impressive, particularly considering that it occurred at a time when the mortality decline in most countries in Asia, Africa, and Latin America was slowing down.[33] The rate of improvement in life expectancy is similar to that experienced by either Costa Rica (at 0.70/year) and Kerala State (at 0.80/year) during their respective periods of rapid mortality decline, but is not as high as Sri Lanka's (at 1.7/year).[2]

Much work on the determinants of reductions in child mortality has focused upon economic factors. It has often been difficult to separate the role of health and other social interventions from general economic progress, since changes have occurred in parallel. In Nicaragua, this was not the case as the sharp drop in child mortality took place at a time when the

TABLE 2—Average Number of Vaccines Administered per Child under the Age of Five

Years	DPT	Polio	BCG
1970–75	0.59	1.02	0.21
1976–80	0.48	1.24	0.17
1981–85	0.99	2.66	0.33

general living standards of the population, measured in terms of per capita national income, actually decreased.

Although there was some improvement in the nutritional status of Nicaraguan children between 1966 and 1988, this was probably more due to a reduced burden of infectious illness than to increased food intake. The factors which led to a drop in the incidence, severity or duration of infection may be the same as those which brought about lower child mortality. It is difficult to follow breastfeeding practice over the last 20 years but it does seem that exclusive breastfeeding is still far from the norm for a variety of reasons.[34] While health education materials were concerned with promoting breastfeeding, the free milk provided by Ministry of Health programs before and after the insurrection may have encouraged mothers to abandon the practice.

The 1974 turning point in child mortality does not appear to have been brought about by a sudden increase in levels of female literacy, although this factor may have contributed to more recent falls in child mortality. Similarly, most of the reduction in mortality due to diseases preventable by vaccination probably occurred some time in the early 1980s and does not coincide with the onset of the decline in child mortality.

The introduction of oral rehydration therapy (ORT) for the management of diarrhea did not occur until the middle of the period of rapidly declining mortality. The Ministry of Health's plan for 1976–80[35] does not even mention ORT, and it was not until 1980 that oral dehydration units were established in health centers and health posts.

The results of this investigation suggest that the availability of water supply and sanitation did not improve significantly from 1974–86. Any improvement in malaria control is unlikely to have been sufficiently great to account for the decline in infant mortality given that malaria is not a major cause of death in this age group in Nicaragua.[36]

The most plausible explanation for the onset in 1974 of Nicaragua's rapid decline in infant mortality can be found in the changes in health services which started to take place at that time. The combined impact of declining per capita numbers of hospitals beds with increasing numbers of health care professionals meant greater availability of personnel to work in the understaffed clinics which had been built in the 1960s. In other words, there was in effect a sharp shift in resources away from hospitals and towards primary and secondary health care. Figure 5 shows how the number of hospital beds per health professional dropped sharply in 1972 and showed a tendency to decrease until at least 1984, parallel to the infant mortality decline.

The reason for this sudden shift in the nature of health care in Nicaragua is complex. The decreasing availability of hospital beds was due to a combination of the devastation resulting from the earthquake, the war, rapid population growth, and economic constraints to public spending. Training of health care professionals, on the other hand, surpassed the increase in population.

Political influences have no doubt also played a part in this situation. The decline and fall of Somoza has been dated from the 1972 earthquake[24,37] and political turmoil has been the norm ever since. A mass popular movement grew up in the late 1970s culminating in the 1979 insurrection. Perhaps sensing that the ground was slipping out from beneath him, Somoza with United States aid began to make significant invest-

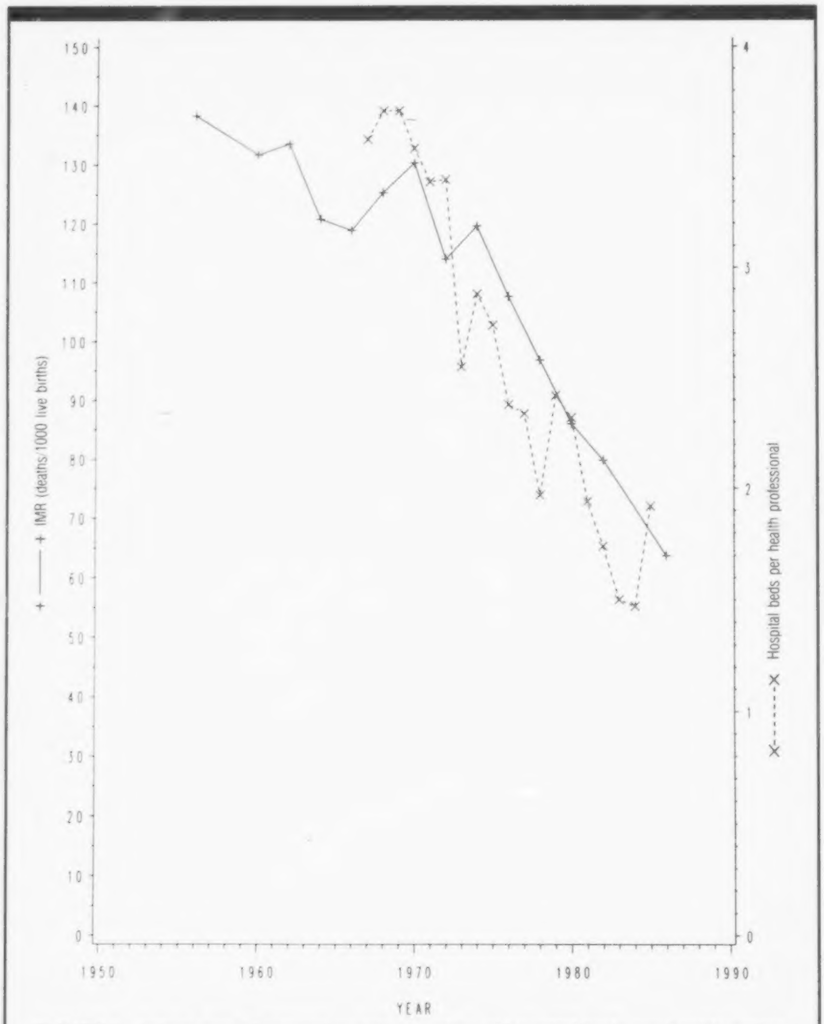


FIGURE 5—Infant mortality trend and the number of hospital beds per health professional in Nicaragua, 1956–86

The solid line is IMR.
The broken line is hospital beds.

ment in social programs. In these programs the government tried to target areas with major guerilla activity.[24]

In post-revolutionary Nicaragua, political factors have been equally important. The Sandinista National Liberation Front took power on a program in which health and education were key components.[38] In President Ortega's speech on the 10th anniversary of the Nicaraguan Revolution, reductions in rates of illiteracy and infant mortality were the first items held up as gains of the revolution.[39]

Cultural factors may also have been important. Although tradition changes too slowly to explain a sudden reduction in mortality of the sort experienced by Nicaragua, certain cultural attributes facilitate rapid reductions in mortality. Primary health care programs can have little impact without appropriate patterns of health-seeking behavior. In Nicaragua, there exists what has been described as a pill culture.[40] People regard orthodox medicine and its providers (i.e. doctors and nurses) with high esteem. Manufactured pharmaceuticals, particularly when given by injection, are generally preferred to traditional herbal remedies although often both will be employed. Nicaraguans readily change treatments when a rapid improvement is not seen and are highly sensitive to the progress of illness in their children. As a result, when health services became more generally available in the 1970s and 1980s, the population quickly made use of them.

It is interesting to compare Nicaragua's period of rapid mortality decline with that of neighboring Costa Rica which also occurred during the early 1970s.[41,42] Both countries had experienced a slow rate of improvement during the 1960s—a phenomenon common to many developing countries.[33] In Costa Rica, the drop in infant mortality was greater in rural areas than in urban areas, both in relative and absolute terms. An analysis of Costa Rica's recent decline in infant mortality also shows up the singular importance of primary health care.[42]

Having reviewed the available data, it is possible to postulate the following scenario. In the late 1960s United States aid was used to pay for the construction of primary health care facilities throughout Nicaragua. In the 1970s, earthquake, war, rapid population growth and economic recession reduced the number of hospital beds per capita. At the same time there was an increase in the supply of nurses and doctors per head of population which made it possible to staff the recently built primary health care facilities. The net effect was an increase in the relative importance of primary and secondary health care vis a vis tertiary health care. The health care facilities provided gained rapid acceptance from the population which already had some cultural affinity for Western medicine.

The social reforms contributing to declining child mortality were made by an isolated and threatened dictatorship in response to the mobilization of impoverished sectors of society. They were continued and extended following the insurrection, as part of the political program of the Sandinista Government which replaced the dictatorship. In early 1990, a new government was elected which is committed to contraction of public expenditure. Monitoring the effects of any change in policy on child health and survival will provide an opportunity to ex-

APPENDIX—Sources of Infant Mortality Statistics in Nicaragua

A review was performed to identify the various sources of data on national child mortality rates in Nicaragua from 1960 to 1989. These are shown in Table A. A note of explanation is necessary about each of the data bases.

The 1971 Census

The last national census of Nicaragua took place on April 20, 1971. Since then, war and lack of funding have prevented the government from carrying out another population census. The 1971 census registered a total population of 1,878,000. The overall rate of omission is thought to be 6.3 percent but in 19 of the 126 municipalities there was an average omission rate of 32 percent and in a further 18 municipalities the rate was 22 percent.[5] An experimental census was carried out in 1969 as a prelude to the 1971 census but its sample was small and unrepresentative of the entire population so the results are not presented here.[43]

The National Demographic Survey of Nicaragua (EDENIC-78)

This was a survey carried out from 1977 to 1978 which included retrospective questions for indirect mortality estimation. It was a follow-up study in which it was intended to compare indirect estimations of mortality with prospective estimates derived from the household change technique. Unfortunately, during the 1979 insurrection all of the prospective data from this survey were lost, leaving only the retrospective questions for mortality estimation. Some of the retrospective data were also lost. The Atlantic Coast Departments of Zelaya and Rio San Juan were excluded from the sampling frame. Although the infant mortality rate in this area is probably higher than in the rest of Nicaragua, it is inhabited by only 8.6 percent of the population.

The Sociodemographic Survey of Nicaragua (ESDENIC-85)

This large household survey took place in 1985, aiming to generate up-to-date demographic and socioeconomic information for the whole country, given that the census planned for 1982 had been suspended. Technical assistance was provided by CELADE and the United Nations Department of Technical Cooperation for Development.[44]

A pilot test in 535 homes was performed in September 1984 and definitive data collection commenced in July 1985. In contrast to EDENIC-78 which excluded the Atlantic Coast, the reference population of ESDENIC-85 covered the entire country. A two-stage cluster sample of 18,000 homes was chosen.

The Sentinel Sites Study

In 1986 the Ministry of Health (MINSA) in collaboration with consultants from UNICEF, established a set of 34 sentinel sites in three regions of the country.[9] These were intended to provide accurate population-based data on morbidity, mortality, and health service utilization for planners and administrators. The sentinel sites were selected by regional teams on the grounds that they be representative of the health conditions of their region. Data were collected by household survey with indirect questions on child mortality.

The National System of Vital Statistics (SINEVI)

In the 1960s, the civil register was administered by the Directorate General of Statistics and Censuses which was part of the Ministry of Economy. In 1970 the DGEC merged with the Department of Economic Studies of the Central Bank to form the Executive Office of Surveys and Censuses (OEDEC). Vital events were registered in the Offices of the Civil Register in each municipality. The data were published annually in the Statistical Yearbook. Coverage was estimated at 70 percent for births and 30 percent for deaths. Only 48 percent of registered deaths were certified by a doctor. The Ministry of Health was responsible for processing cause-specific mortality rates which were not published in Nicaragua but were sent to the World Health Organization.[45]

When the new government took power in 1979, several changes were introduced to improve the registration of vital statistics but, despite these measures, serious deficiencies in death registration remain. In 1987, overall coverage was estimated at 47 percent,[36] but only about 34 percent of infant deaths are reported.[11] Birth registration is somewhat better; in 1983 coverage was calculated to be somewhere between 79.3 and 86.5 percent, varying from less than 60 percent in the rural population to about 95 percent in the urban population.[46]

TABLE A—Sources of Infant Mortality Statistics in Nicaragua

Date	Data Source	Type of Data	No. Women Ages 15–34	Regional Coverage[a] %
1971	National census[6,7]	Population census	256,907	100
1978	EDENIC-78[8]	Sample survey	6,237	91
1985	ESDENIC-85[9]	Sample survey	22,000[b]	100
1986–88	Sentinel sites[10]	Sentinel sites survey	9,615	54
1956–86	SINEVI[11]	Death registration	—	100
1986	SINEVI-86[12]	Birth registration	112,200[b]	100

Source: References 6–12.

[a] The proportion of the inhabitants of the country included in the reference population.

[b] Approximate only.

plore the validity of the explanations advanced in this paper. □

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The Role of Attitudes, Beliefs, and Personal Characteristics of Italian Physicians in the Surgical Treatment of Early Breast Cancer

ABSTRACT

The influence of Italian physicians' attitudes, beliefs, and personal characteristics on medical decision making is examined in the case of surgical treatment of early breast cancer. Responses to a mail survey of 657 physicians from different specialties were analyzed comparing doctors recommending a radical procedure (9%) to those preferring a conservative procedure for younger patients only (25%), and those considering conservative surgery the treatment of choice regardless of patients' age (66%). The findings suggest that the likelihood of physicians' preferring a conservative procedure is influenced by their specialty and the extent to which they feel that a patient should have a role in the treatment decision more than by differences in the beliefs of treatment outcomes. Only preferences of the small group indicating radical surgery as the sole admissible treatment can be accounted for by ignorance or distrust of results of recent trials. These findings suggest that other than scientific factors guide many doctors in their decision making; they may help to explain why the diffusion of research results into clinical practice is often disappointingly slow. (*Am J Public Health* 1991; 81:38–42)

Alessandro Liberati, MD, Giovanni Apolone, MD, Antonio Nicolucci, MD, Carlo Confalonieri, MD, Roldano Fossati, MD, Roberto Grilli, MD, Valter Torri, MD, Paola Mosconi, BiolSci, and Alessandro Alexanian, MD

Introduction

The extent to which trial results influence clinical practice is a subject of growing interest and controversy. The roles of doctors' attitudes, belief and personal factors in this respect have seldom been directly assessed, however. The controversy on the optimal surgical treatment of early breast cancer is an excellent area for such an investigation. Despite the substantial body of evidence now available on the comparative efficacy of limited and radical surgery,[1–3] the acceptance of conservative procedures in general practice is slower than one could expect. Possible explanations offered include: ignorance of trial results; uncertainty in interpreting trial results[4–9]; the influence of personal or institutional-related factors.[10–12]

To shed light on the merits of these three related hypotheses, we undertook this study as part of a current research program aimed at assessing quality of care for breast cancer patients in general hospitals.[6]

Organization of Cancer Care in Italy

In Italy health care is provided largely free of charge with patients co-payment for a limited—though increasing—number of services. The national health service—launched in 1978—envisages a system where health care is provided through public facilities (approximately 1,000 general hospitals where about 40,000 doctors practice) or, indirectly, by contracting with private health providers. Relative to cancer care apart from patients seen at the few Italian cancer centers, most patients are cared for at public general hospitals. There is a great deal of variability in the type of existing organization within these hospitals: few are equipped with oncology departments staffed with board-certified oncologists; others—accounting for approximately 10 percent of the total—see patients in oncologic wards staffed by internists, surgeons, and radiotherapists with special interest in oncology; but the vast majority of hospitals have no ad hoc organization and patients are seen at different departments depending on their specific needs.

Methods

This study was conducted between October 1986–June 1987 on a sample of hospital-based physicians. Sixty hospitals taking part in an existing monitoring program on quality of cancer care were invited to participate, 65 percent agreed to do so. Participating and non-participating hospitals were similar in size (median number of beds of 888 and 870, respectively), and organization of cancer care (35 percent of both groups had radiotherapy on site and an oncology department was present in 20 percent and 30 percent of participating and non-participating cen-

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From the Laboratorio di Epidemiologia Clinica, Istituto di Ricerche Farmacologiche, Mario Negri, Milano, Italia (Liberati, Apolone, Nicolucci, Fossati, Grilli, Torri, Mosconi, Alexanian), and the Divisione di Medicina, Ospedale Civile di Rho, Rho, Italia (Confalonieri). Address reprint requests to Dr. Alessandro Liberati, Laboratorio di Epidemiologia Clinica, Centro di Coordinamento GIVIO, Istituto di Ricerche Farmacologiche Mario Negri, Via Eritréa 62, 20157, Milano, Italia. This paper, submitted to the Journal November 9, 1989, was revised and accepted for publication April 6, 1990.

ters, respectively). Six hundred and fifty-seven questionnaires were returned corresponding to an overall response rate of 68 percent (657/970), 20 percent of which were from physicians not seeing patients with breast cancer. Surgeons had a 60 percent response; internists 68 percent, gynecologists 74 percent, radiotherapists 84 percent and oncologists 89 percent. Participation tended to be slightly higher among younger (78 percent) and female (77 percent) physicians. An additional telephone inquiry was attempted on a randomly selected one-third of the 313 nonresponders. All were successfully traced and their responses were similar to those obtained by mail.

In the questionnaire, basic data were requested about the hospital in which physicians practiced and some personal characteristics. The first section of the questionnaire described four scenarios, two dealing with treatment of primary breast lesion and two dealing with adjuvant postsurgical treatment. In this paper we shall discuss the results relative to the surgical preferences. The first scenario was phrased as follows:

> A 35-year old premenopausal patient presents with a 1.5 cm. diameter mass in the upper outer quadrant of her right breast. Axillary lymph nodes are clinically negative and there are no other signs of metastases. Assuming that whatever blood test, X-ray examinations, and nuclear scans you use are negative and that a preliminary biopsy has established the diagnosis of carcinoma, do you favor a radical surgical approach or a breast-sparing operation (limited surgery) combined with primary radiotherapy? The radical surgical approach would mean a type of radical mastectomy (i.e. Halsted or Patey), while the limited surgery would involve a quadrantectomy, plus axillary dissection and radical radiotherapy.

After checking one of the two options (1 = I favor a radical surgery approach; 2 = I favor a limited surgery approach), responders rated the superiority of the alternative treatment modes in achieving the following outcomes: local control, disease-free survival, avoidance of morbidity, good cosmetic results, and long-term survival. A 5-point scale was used for each outcome: 1 = limited surgery plus irradiation is vastly superior to radical surgery; 5 = radical surgery is vastly superior to limited surgery plus irradiation; 3 = the two treatment modes are equally effective.

A second case presentation referred to an older patient (i.e. 60 years) with identical clinical information to see whether doctors deemed patient age a critical factor in treatment indication. The same treatment choices and efficacy judgment were required.

TABLE 1—Distribution of Physicians' Sociodemographic and Practice-related Characteristics according to Therapeutic Preferences

	R N = 33 (9%)		C-R N = 112 (25%)		C N = 298 (66%)	
Demographics	n	(%)	n	(%)	n	(%)
Age (years)						
≤35	9	(27)	37	(33)	120	(40)
36–50	18	(55)	56	(50)	139	(47)
>50	6	(18)	19	(17)	39	(13)
Sex						
Male	29	(88)	104	(93)	260	(87)
Female	4	(12)	8	(7)	38	(13)
Hospital Size (beds)						
≤400	4	(12)	23	(20)	65	(22)
401–1000	15	(46)	50	(45)	113	(38)
>1000	14	(42)	39	(35)	120	(40)
Patient Volume**						
<20	9	(28)	30	(27)	68	(23)
20–50	13	(41)	48	(44)	106	(36)
>50	10	(31)	32	(29)	118	(41)
Availability In-house Radiotherapy Equipment						
No	14	(42)	44	(39)	106	(36)
Yes	19	(58)	68	(61)	192	(64)
Specialty* (p = 0.0003)						
Internal medicine	6	(9)	22	(31)	42	(60)
Surgery	23	(11)	61	(29)	127	(60)
Radiotherapy	2	(3)	3	(5)	54	(92)
Medical oncology	1	(2)	12	(21)	43	(77)
Gynecology	1	(2)	12	(27)	31	(71)

R = radical surgery regardless of patient's age;
C-R = conservative only in younger patients;
C = conservative surgery regardless of patient's age.
*Information not available in three cases.
**Information not available in nine cases.

In the second section, physicians were asked to comment on findings from a pattern of care study conducted in Italy in 1984 documenting that only approximately one out of four patients eligible for limited surgery actually had limited surgery.[6] Doctors were asked to report why in their opinion a radical procedure was still the preferred treatment option by choosing among a list of possible explanations reported into the questionnaire. A free text format was also included so that they could add other reasons if they were not satisfied by the list.

The final section of the questionnaire contained a series of statements designed to measure physicians' attitudes toward involving patients in treatment decisions. Responders indicated on 5-point scales the extent to which they agreed or disagreed with 10 phrases such as "asking patients to participate in treatment decisions produces unnecessary stress for them," or "patients who participate in treatment decisions are less anxious and depressed during recovery." Statements were worded so that, for some, agreement indicated attitudes favoring participation and, for others, an unfavorable attitude. The scale was similar to the one used in a previous study in the United States.[10]

Analysis of the role of physicians' willingness to involve patients in treatment decisions was done both by considering answers to individual statements and by analyzing the overall score based on the scale; the Kruskal-Wallis test was used for statistical analysis.[13] In the computation of this test, all of the scores from all samples combined are ranked. The sum of the ranks in each sample is found, and the test determines whether these sums of ranks are so disparate that they are not likely to have come from samples which were all drawn from the same population.[13] Results of this type of analysis are those reported in Table 2.

Results

Table 1 shows the distribution of physicians' sociodemographic and practice-related characteristics according to therapeutic preferences. Two-thirds felt that a conservative procedure was the appropriate treatment independently of patients' age; only 33 (9 percent) preferred the traditional radical procedure regardless of age. Almost all the radiotherapists favored the conservative procedure regardless of patients' age while more than one-fourth of doctors from other specialties deemed the conservative procedure appropriate only for younger patients. Surgeons and internists were more often supportive of the traditional radical technique.

A greater willingness to involve patients in the decision-making process was associated with preference for conservative surgery (Kruskal-Wallis test = p = 0.011 for difference in overall patient participation score among the three groups). Detailed analysis of physicians' answers to individual statements component of the overall score indicated that doctors recommending conservative surgery regardless of patient age (the "C" group in the tables) had a consistently different pattern of attitudes (see Table 2) compared to those indicating radical surgery as the optimal treatment.

Besides this marked difference in personal attitudes, the groups also dif[illegible] in their perceptions of the efficacy [illegible] and overall desirability of the two trea[illegible]ments (Table 3). Almost all doctors in favor of the more radical procedure (94 percent) thought it was superior in terms of local control, particularly for the older woman while doctors always favoring limited surgery most of time claimed comparable effectiveness of the two treatments in all but cosmetic results.

In the intermediate group (i.e. those preferring conservative surgery in the younger but radical in the older patient) the relationship between reported therapeutic preference on one side, and attitudes and beliefs about outcomes on the other, was less easy to interpret. Patients' age seemed, although to a limited extent, to influence perception of treatment effectiveness in this group. Physicians' answers were more like those of the radical

TABLE 2—Distribution of Scores (mean ranks) in Individual Statements Exploring Physicians' Willingness to Involve Patients in Treatment Decisions according to Therapeutic Preferences

Negatively worded statements	R	C-R	C	p value
	Mean Ranks			
Patients may lose confidence in their physician if they believe that he/she has no firm opinion about the best treatment	221.1	227.9	219.8	.80
Encouraging patients to participate may do more harm than good	197.6	210.5	229.0	.15
Patients can't possibly make good decisions because they don't understand information	215.1	200.2	230.9	.05
Asking patients to participate in treatment decisions produces unnecessary stress	198.3	194.6	234.9	.003
Even if they receive enough information most patients are too upset to make a decision	216.8	197.5	231.8	.03
Positively worded statements	**R**	**C-R**	**C**	**p value**
Patients who participate in treatment decisions make a better adjustment to the disease	225.1	253.6	209.8	.003
Patients should have a greater influence on treatment decisions than their doctor	243.2	241.8	212.2	.05
Patients who participate in treatment decisions are less anxious and depressed	245.3	244.4	210.9	.02
If given comprehensible medical information pts. can make good decisions about treatment	248.9	249.4	208.6	.004
Most patients want to be involved in treatment decisions	217.8	249.2	212.2	.02

R = radical surgery regardless of patient's age;
C-R = conservative only in younger patients;
C = conservative surgery regardless of patient's age.
Significance tests use Kruskall-Wallis One-Way Analysis of Variance.[13] See text for further information.

TABLE 3—Proportions of Doctors with Different Beliefs about the Yield of Radical and Conservative Surgery according to Specific Treatment Outcomes

		Therapeutic preferences					
		R		C-R		C	
Type of Treatment Outcome	Perceived efficacy	35y old	60y old	35y old	60y old	35y old	60y old
Prevention of local recurrence	C>R	6	3	10	10	8	7
	C=R	9	3	62	42	77	76
	R>C	84	94	28	48	15	17
Disease-free interval	C>R	—	—	4	4	4	4
	C=R	18	9	83	63	93	93
	R>C	82	91	13	33	3	3
Risk of morbidity	C>R	6	6	28	21	24	25
	C=R	24	24	55	49	59	59
	R>C	70	70	17	30	17	16
Cosmetic results	C>R	94	76	95	84	96	94
	C=R	3	19	3	13	2	4
	R>C	3	5	2	3	2	2
Long-term survival	C>R	3	3	6	4	4	4
	C=R	24	12	82	63	94	93
	R>C	73	85	12	33	2	2

R = radical surgery regardless of patient's age;
C-R = conservative only in younger patients;
C = conservative surgery regardless of patient's age.

Legend:
C>R = Doctors considering conservative surgery superior to radical;
C=R = Doctors that attribute the same effectiveness to the two treatments;
R>C = Doctors considering radical surgery superior to conservative.

TABLE 4—Physicians' Explanations of the Limited Popularity and Preference for Execution of Conservative Surgery in Younger Patients Documented in a Nationwide Pattern of Care Study in 1984 (Italy) (see ref. 6)

	R		C-R		C	
	n	(%)	n	(%)	n	(%)
Trial results still controversial	14	(42)	53	(47)	53	(18)
Technical difficulty of cons. surgery recommends its performance mostly in younger patients	7	(21)	20	(18)	21	(7)
Cons. surg. is limited by the availability of radiotherapy equipment	10	(30)	40	(36)	122	(41)
Surgeons are still reluctant to accept trial results	2	(6)	11	(10)	101	(34)
Cosmesis not yet sufficiently considered	2	(6)	7	(6)	25	(8)
Others	2	(6)	7	(6)	26	(9)

R = radical surgery regardless of patient's age;
C-R = conservative only in younger patients;
C = conservative surgery regardless of patient's age.
Column percentages exceed 100 because doctors could give more than one answer.
x^2_{10} d.f. = 25.6 $p < 0.03$

group particularly with respect to the decision pattern for the older patient (Table 3).

As a whole, joint examination of the beliefs of the intermediate group (Table 3) and their willingness to involve patients in treatment decisions (Table 2) suggests that—independently of perceived treatment effectiveness—less consideration for patients' expectations was probably the factor influencing choice of a radical procedure for an older patient in whom negative cosmetic consequences were judged less important. This interpretation is supported indirectly considering the distribution of physicians' opinions of reasons why the diffusion of conservative surgery is still limited (Table 4). The majority in the radical group (42 percent) claimed conflicting trial results as the reason for this, but 34 percent of doctors in the conservative group felt that prevailing practice could be accounted for by surgeons' reluctance to accept trial results (Table 4).

Discussion

Decisions about alternative treatments involve implicit and explicit weighing of trade-offs among different outcomes (survival, morbidity, functional status, etc.). Our study suggests that beliefs about treatment outcomes derived from clinical trials are not always the most important factors in decisions for some doctors. This is not a new finding as other studies exploring determinants of physicians' practice[10–14] have found associations of specialty, age, and practice-related characteristics with practice patterns.

The grouping of doctors according to their therapeutic preferences is of particular interest for analyzing the consistency between perceived outcomes of treatment and the decision pattern. At one extreme, we found a small group (9 percent) whose therapeutic preference—although based on a miscomprehension, or non-acceptance, of recent scientific literature—was consistent with perceived outcomes of treatment: doctors chose a radical procedure because they thought it was more effective than limited surgery. If we take their answers as a true reflection of their opinions, and not as the result of a preference-dependent selection of that segment of the literature fitting their own preconceptions, exposure to the relevant scientific literature would be the appropriate remedial action.

At the other extreme, those choosing limited surgery appeared to be guided by the consideration that if this approach is just as effective as a radical procedure, the less mutilating treatment should be preferred.

Although the beliefs of the "intermediate" (i.e. radical-conservative) group about treatment effectiveness were similar to those of the "conservative group," radical surgery was still their preferred treatment for a 60 year old patient. While we cannot rule out that this was due to some anticipated contraindication for limited surgery in the older patient, their lack of consideration for the patient's right to share in the decision-making appears a more convincing explanation of their different behavior. When facing the same scientific uncertainty less "paternalistic" doctors (i.e. those attaching greater importance to patients' opinions) appear to be influenced by patient-centered outcomes (i.e. cosmetic) more than those who think they know what is in their patients best interest, thus relying more on physician-centered end-points (i.e. disease free-survival, total survival).

The original sample was drawn from among physicians working at a group of general hospitals participating in a program of quality of care evaluation.[6] Even if for some reason participating hospitals were not fully representative of all Italian physicians, this does not affect the internal validity of our findings. Furthermore, our findings favorably compare with other surveys on specialists[14] and non-specialists[10,17,18] in different countries. Studies from the UK,[18] USA,[10] and Canada[14] all consistently showed that surgeons, and to a limited extent also internists not specialized in oncology, are still more often in favor of radical procedures.

Data on actual practice would be stronger evidence of treatment preferences, but they are seldom available. Given that our study was done in a group of hospitals where a monitoring program on care given to breast cancer patients was going on,[6] it may be interesting to look at some comparisons between the two sources of information. Consistent with the key role of surgeons' opinions in the decision about surgery, we found—within a group of more than 1,000 cases of newly diagnosed patients in 1984–85[6]—that limited surgery was offered only to 25 percent of eligible patients. Also, data on actual practice confirmed the age-dependent orientation of doctors in that the more conservative operation was given to 34 percent of women younger than 50 years and compared to 21 percent among older patients.[6]

The public health implications of our study depend on the path of therapeutic decisions, which in turn depend on health care organization in different countries. In Italy, where cancer care is still seldom interdisciplinary, especially in community hospitals, it is worrisome that as many as 40 percent of surgeons—i.e. the specialists who decide about surgery—still favor radical surgery for an older patient and 11 percent even for a younger patient. This should not be seen as totally surprising since at the time this survey was carried out official recommendations of the Italian Breast Cancer Task Force were still in a transition period. They left open the choice of the most appropriate treatment indicating as acceptable both limited surgery and a more radical procedure (i.e. the Patey type of "modified" radical mastectomy).

Where cancer care is multidisciplinary (like in the US or Canada), it is likely that the more modern opinion held by other specialists is already forcing surgeons to change opinion and consequently practice. Less easy to predict is the extent to which changes in doctors' perception of patients' rights in treatment decision will affect prevailing practice in the near future. Considerations of this sort should be kept in mind when comparing results of patterns of care studies and physicians' surveys. ☐

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Maternal Feeding Behavior and Child Acceptance of Food during Diarrhea, Convalescence, and Health in the Central Sierra of Peru

Margaret E. Bentley, PhD, Rebecca Y. Stallings, MHS, Mary Fukumoto, PhD, and John A. Elder, PhD

ABSTRACT

Feeding patterns by mothers and child acceptance of food were measured in a Peruvian village to determine changes on days when children had diarrhea as compared to days of convalescence and health. Morbidity surveillance identified 40 children, aged 4–36 months, with diarrhea. Children were followed using twelve-hour in-home structured observations during two to four days each of diarrhea, convalescence, and health. Using scales of maternal encouragement to eat and child acceptance of food and cumulative logistic regression analyses, maternal encouragement to eat decreased significantly during convalescence compared to diarrheal days (OR: 0.54, 90% CI: 0.35, 0.82) and health compared to diarrhea (OR: 0.65, 90% CI: 0.46, 0.93). In contrast, child acceptance of food increased during health compared to diarrhea (OR: 1.55, 90% CI: 1.02, 2.35). Results illustrate the importance of carefully examining the behavioral aspects of nutritional intake. Decreases in intake during diarrhea are due to anorexia and not withdrawal of food by mothers. In response to reductions in child appetite during illness, mothers are more likely to encourage children to eat, while they tend to become more passive feeders after the diarrhea has stopped. Program efforts should focus on messages to feed children more actively especially after diarrhea episodes, when appetite levels increase. (*Am J Public Health* 1991; 81:43–47)

Introduction

Worldwide studies of infant and child nutrition have demonstrated that poor infant feeding and weaning practices along with frequent infections, in particular diarrheal diseases, are important determinants of growth faltering and malnutrition in the developing world.[1–3] However, the primary focus of diarrheal disease control (CDD) programs has been on improved case management through the promotion of oral rehydration therapy, while dietary management has been relatively neglected. Where program messages do exist, they are generally quite vague—such as "continue feeding during diarrhea"—and therefore may be difficult for mothers and other caretakers to implement.[4]

The development of specific recommendations for feeding practices requires information about existing infant feeding practices during illness. In the absence of data, the public health community has often assumed that mothers in nearly all cultures withhold food from their children during diarrhea episodes.[5–8] This implies that modification of this ill-advised presumed behavior will reduce the negative nutritional effects of diarrhea. Such a perspective ignores the possibility that mothers may already be feeding their children during diarrhea, and that diarrhea-associated anorexia may be the major determinant of reduced dietary intakes.[9–10]

This study focused on the behavioral factors related to feeding during diarrhea, and assessed the relationship of illness to appetite in young children. In particular, we quantified the degree of effort caretakers applied to encourage children to eat and the acceptance of food by children during diarrhea episodes, the immediate postconvalescent period, and subsequent, symptom-free days of "health."

Methods

Data Collection

The study was conducted in a rural Andean village in the Callejon de Huaylas, Peru as part of the Dietary Management of Diarrhea program, funded by the United States Agency for International Development, Office of Nutrition. It was implemented in Peru by the Institute of Nutrition Research and the Federal and State Ministries of Health in collaboration with Johns Hopkins University. The project has been described in greater detail elsewhere.[11–12]

Morbidity surveillance in the study area identified children with diarrhea by visiting each household at 48-hour intervals. A total of 118 children, between the ages of 4–36 months, were under surveillance for 12 months. Diarrhea was defined as four or more liquid or semi-liquid stools/day, convalescence as 0–1 liquid or semi-liquid stools/day immediately after at least two days of diarrhea, and healthy observation days as an absence of any illness symptoms. This latter requirement was quite problematic, since children in this setting were frequently ill with some sort of infection.

From the Division of Human Nutrition, Department of International Health, Johns Hopkins University (Bentley, Stallings, Elder) and Instituto de Investigacion Nutricional, Lima, Peru (Fukumoto). Address reprint requests to Margaret E. Bentley, PhD, Division of Human Nutrition, Department of International Health, Johns Hopkins University, 615 North Wolfe Street, Baltimore, MD 21205. This paper, submitted to the Journal July 21, 1989, was revised and accepted for publication May 3, 1990.

TABLE 1—Mean Energy (Kcal/kg/day) and Protein (g/kg/day) Intake by Stage of Illness among Breastfed Children 4–36 Months of Age

Intake	Stage of Illness		
	Diarrhea	Convalescence	Symptom-Free
	n = 36	n = 36	n = 36
Kcal/kg/day	76*	85	86*
g/kg/day	1.3†	1.5	1.6

*Difference = 10 kcal/kg/day (p = 0.052)
†Difference = 0.3 g prot/kg/day (p = 0.009)
SOURCE: Esrey *et al*, FASEB 1988

Twelve-hour in-home structured observations of maternal and child feeding behavior were completed, during two to four days each of diarrhea, convalescence, and health. Forty children (drawn from the population under surveillance) were studied for the present study. Children were eligible for this study if, on the day of morbidity surveillance they met the criteria for diarrhea, if the episode was no more than one or two days old, if the children were not exclusively breastfed, and if a data collector was available to begin a set of 12-hour observations. Children were selected for across a full calendar year.

Weighed dietary intakes of the same children, reported in detail elsewhere,[13] demonstrate that there is a decrease in dietary intakes during diarrhea (Table 1). What remains to be explained is why these reductions occur.

A data sheet was developed to code observations of maternal feeding behaviors and child acceptance of food during each meal. A total of 787 meals were recorded for the 40 children. For 75 percent of the meals, mothers were the feeders, so we refer to "maternal feeding behavior" in our discussion. Neither the identity of the caretaker nor that of the observer was found to be significantly associated with either the mother's or child's behavior (data available from authors on request).

Behavioral observations were made at the level of the food or drink, for a total of 951 observations. Feeding observations of commercial oral rehydration solution (ORS) and non-caloric teas have been excluded. If a child consumed potatoes and soup during the same meal, the data collector would code child acceptance of food behaviors separately for each food. This was done because field testing showed that behaviors varied within a single meal episode, depending on the food offered. In addition, we wanted to assess whether children preferred or rejected specific types of food, either during illness, recovery, or health.

Data collectors were aware only that the purpose of the research was to measure diet and feeding behavior during illness, along with other household treatment behaviors. To check whether the invasive nature of the study methodology (that placed data collectors in homes for repeated 12-hour observations) may have influenced maternal feeding behavior, a series of three-hour feeding observations without weighing either food or children before and after breastfeeding were completed for 30 children. There were no significant differences in the frequency or duration of breastfeeding between the three-hour and 12-hour observations (data available from author on request).

Scale Construction

The data sheet contained five items measuring maternal feeding behavior and four items measuring child acceptance of food. Because these items had a logical order (i.e. mothers would not physically force children to eat if they had not already encouraged and pressured them verbally to eat), Guttman scaling techniques were used for scale construction.[14–15] The coefficients of reproducibility were quite high, .98 for Maternal Encouragement to Eat and .97 for Child Acceptance of Food, indicating that the observed response patterns fit closely the response patterns predicted by the Guttman analysis. After testing, three items remained in each scale.

For the Maternal Feeding Behavior scale the final items were: 'Verbally Encourage'→ 'Verbally Pressure'→ 'Physically force.' For the Child Acceptance of Food scale the items were: 'Rejects Food'→ 'Appetite Level'→ 'Asks for Food.' These two scales were then used as the dependent variables in the statistical analyses.

The Child Acceptance of Food scale was validated against actual intakes by comparing the mean scale for the day to the number of kilocalories from non-breastmilk for the day. Because the scale is ordinal rather than interval, Spearman's rho formula was used to calculate the correlation. The final Spearman correlation coefficient was .47, which shows a strong relationship between the scale and actual intakes.

There is no obvious means of validating the Maternal Encouragement to Eat scale against actual intakes. Increased encouragement to eat is a response to decreased appetite. The lower the appetite, the more encouragement children receive. Since maternal encouragement may or may not be successful, it is not clear that actual intakes should be associated with encouragement.

Cumulative Logistic Regression

To preserve the ordered categorical nature of the scales, cumulative logit procedures were used[16] and separate logistic regressions were performed for each cut-point (i.e. each ascending value on the scale). Because our early models indicated that the magnitude and direction of the beta estimates for various factors were not consistent across these cut-points, we were not able to make an assumption of "proportional odds," which would have allowed us to report only one set of final beta estimates. For the Maternal Feeding Behavior scale, only two cut-points were used because just 13 observations scored 3, which was too small a percentage of observations to compare with the 0–2 category.

The main hypothesis was that "stage of illness" affects both maternal and child feeding behavior, so this variable was entered into the model first. Then, because we knew that age should theoretically influence feeding behavior, we added age to the model as a categorical variable. The study design followed children prospectively through diarrhea, convalescence, and health, as they were aging. Children 4–5 months old were omitted from these analyses because they were few in number and inconsistent in preliminary analyses. Even with age in the model, "stage of illness" was significantly associated with feeding behavior and child acceptance of food.

Two different techniques were used to test for possible confounding factors, one following Kleinbaum, *et al*,[17] and one following Greenland.[18] Both were modified to allow the examination of polychotomous dependent and independent variables. On the Maternal Feeding Behavior

TABLE 2—Final Logistic Regression Model, Mother's Encouragement to Eat Score (1st Maternal Feeding Behavior Scale Cut-Point: 0 vs 1–3)

	Odds Ratio	90% Confidence Interval
Intercept	0.66	(0.36, 1.24)
Stage of illness		
Ref: Diarrhea	1.00	—
Convalescence	0.54	(0.35, 0.82)
Health	0.65	(0.46, 0.93)
Age (months)		
6–11	0.82	(0.35, 1.93)
12–17	1.65	(0.70, 3.88)
Ref: 18–36	1.00	—
Consistency		
Ref: Liquids/semi-solids	1.00	—
Solids	0.39	(0.30, 0.51)
Breast milk consumed (kcal) during 12-hour observation		
Ref: None	1.00	—
≤410	1.82	(0.87, 3.80)
>410	1.63	(0.69. 3.84)

TABLE 3—Final Logistic Regression Model, Child's Acceptance of Food Score (2nd Child Food Acceptance Scale Cut-Point: 0,1 vs 2,3)

	Odds Ratio	90% Confidence Interval
Intercept	1.50	(0.81, 2.76)
Stage of illness		
Ref: Diarrhea	1.00	—
Convalescence	1.12	(0.77, 1.64)
Health	1.55	(1.02, 2.35)
Age (months)		
6–11	1.49	(0.92, 2.40)
12–17	0.90	(0.55, 1.48)
Ref: 18–36	1.00	—
Breast milk consumed (kcal) during 12-hour observation		
Ref: None	1.00	—
≤410	0.48	(0.31, 0.74)
>410	0.28	(0.15, 0.49)
Consistency		
Ref: Liquids/semi-solids	1.00	—
Solids	0.62	(0.47, 0.82)
Kilocalories consumed at the last feeding, either breastmilk or solid food		
Ref: 0–33.4	1.00	
33.4< <64	0.58	(0.43, 0.77)
≥64	3.32	(2.42, 4.56)

scale, variables were retained in the model if they were found to be confounders at either cut-point, while on the Child Acceptance of Food scale, variables were retained if they were found to be confounders at two of the three cut-points. Variables tested include, among others, type and consistency of food, caloric density of the food, kilocalories eaten at last meal, time since last meal, amount of breastmilk consumed during the observation period, the number of breastfeeding episodes, time of day, number of meals during the day, sex of child, identity of caretaker, and number of foods at a meal.

The variable fever (measured with a fever strip and later dichotomized) was deleted from the final analyses because, although highly significant in the preliminary analysis, the number of observations when fever was present (n = 17) was too small for multivariate analyses. However, a simple odds ratio of fever with appetite scores at cut-point 2 (corrected for intra-child correlation) shows that fever is much more likely to be associated with lower appetite (odds ratio = 0.06, 90% confidence interval: 0.01 to 0.46).

The final models were developed by adding to the model of stage of illness and age those variables that proved to be confounders. In addition, food consistency was added to both models since it was found to be closely related to both child appetite and maternal encouragement. Final variance estimates were corrected for the presence of intra-child correlation due to the longitudinal nature of the study design.[19]

Results

In the interest of brevity, the final models presented are those with cut-points closest to the median for both scales (Tables 2 and 3). However, when results for other cut-points differ significantly for the variables of primary interest, these are discussed.

For the Maternal Feeding Behavior scale (Table 2), when controlling for age and the other factors shown, the scores were significantly lower during convalescence and health compared to diarrhea (for the second cut-point, i.e., 0–1 vs 2–3, scores were not significantly lower for convalescence). In other words, mothers were more likely to encourage their children to eat during diarrhea episodes, compared to the other two intervals. In addition, mothers gave more encouragement when their children were eating liquid or semi-solid foods.

For the Child Acceptance of Food model (Table 3), when controlling for age and the other factors shown, children had lower food acceptance scores during diarrhea compared to health. For this cut-point, food acceptance scores were not significantly higher during convalescence, compared to diarrhea, but they were significantly higher in the model for the third cut-point (i.e., 02 vs 3).

Acceptance of food scores were higher when children were presented liquid or semi-liquid foods, compared to solids, regardless of stage of health. Twelve-hour breast milk intake was negatively associated with appetite on the child scale, indicating that higher levels of breast milk intake reduced appetite for non-breast milk foods. Breast milk consumption was measured by weighing the children immediately before and after breastfeeding.

Interestingly, in the preliminary screening, Kcals from breast milk consumed between meals was not signifi-

TABLE 4—Cross-Tabulation of Maternal Encouragement to Eat Scale and Child Acceptance of Food Scale

Maternal Scale (Low to High)	Count Row% Col%	Child Scale (Low to High)				
		0	1	2	3	
	0	12	191	254	79	536
		2.2	35.6	47.4	14.7	
		27.9	63.5	67.6	70.5	
	1	13	91	112	31	247
		5.3	36.8	45.3	12.6	
		30.2	30.2	29.8	27.7	
	2	18	19	10	2	49
		36.7	38.8	20.4	4.1	
		41.9	6.3	2.7	1.8	
		43	301	376	112	832

Goodman and Kruskal's Gamma = −.23851

cantly associated with the child's food acceptance at the subsequent meal, although the coefficients for this factor were negative, as was the case for 12-hour breast milk intake. We interpret this to mean that breastfeeding is associated with an overall reduction of child appetite for other foods, which cannot be adequately assessed at a meal level. This relationship did not appear to be confounded by time, as time since last breastfeeding was not found to be significantly associated with child's food acceptance.

Another function included in the screening, namely Kcal from either breast milk or non-breast milk consumed during the last feeding episode, was positively associated with child appetite at the current meal when comparing a high level of consumption (≥64 Kcal) to a low level (≤33.4 Kcal). However, the relationship was negatively associated when comparing an intermediate level (between 33.4 and 64 Kcal) to a low level. This same inconsistency was observed for the first cut-point, but for the third cut-point, both the intermediate and high levels were positively associated with child appetite. Because the positive association is consistent with the preliminary analyses, we expect that this is the correct direction.

A cross-tabulation of the scores on the Maternal Encouragement to Eat Scale and those on the Child Acceptance of Food Scale (Table 4) shows a clear negative relationship between the two scales, suggesting that mothers were increasing their efforts to feed their child as the child's appetite decreased. This relationship would have appeared stronger if the effect of food type was removed from both scales because the logistic analysis showed that both child acceptance and maternal encouragement increased when liquid and semi-solid foods were involved. The table also shows that these mothers were generally passive feeders, 64 percent of the time giving no encouragement to eat. These data corroborate the same finding in our early ethnographic work on this population. However, at the lowest level of child acceptance of food, over 70 percent of the time, mothers gave some encouragement.

Discussion

The results from this study demonstrate that these Peruvian mothers, normally quite passive child feeders, were more likely to encourage their children to eat during diarrhea episodes compared to both convalescence and health. Children experiencing diarrhea did appear to have observable reductions in appetite, and more often rejected food. These results corroborate our ethnographic data. Mothers frequently voiced concern that their children were "weak" and had reduced appetites during diarrhea.[11]

The more total kilocalories consumed by the child during the previous meal or breastfeeding, the better the appetite at the current meal. At first, this seemed counter-intuitive, but Esrey, *et al*,[13] found that these children averaged only 71 percent to 91 percent of recommended energy intakes, so if the children were not suffering from diarrhea or fever-induced anorexia, they would probably be willing to eat whenever food was offered. If this was the case, a child's appetite at any given meal would reflect the general level of health for the day rather than the amount consumed at the previous meal/breastfeeding, which was borne out by the positive relationship between appetite at the current meal and kilocalories consumed at the previous meal/breastfeeding. Again, this relationship did not seem to be confounded by time since last meal/breastfeeding, which was tested and was not significant.

What is the public health significance of these data? First, it is clear that in this setting mothers are not withholding food from their children when they experience diarrhea. Where this issue has been closely studied in other sites, similar results have been found. For example, in Nigeria, where we have used the same methodology as in this Peruvian site, mothers appear to force-feed their children more often when they have diarrhea, possibly as a response to their children's refusal to eat.[20] In India, some mothers shift the diet during diarrhea toward foods that they perceive as "helpful" and away from foods perceived as "harmful."[21] In each of these three sites, mothers' concern about appetite loss during diarrhea was high.[11,20-22] We hope that these data will replace the anecdotal information and survey data that reports that mothers "starve" their children during diarrhea episodes.

The role of anorexia during diarrhea appears to be an important factor in reductions of dietary intake during illness. We are not aware of other studies that have attempted to measure the degree of appetite of children in a field setting, although one clinic-based study suggested that anorexia was an important mechanism for reduced dietary intake during diarrhea.[23] Although it is expected that children's appetites should rise during the convalescent phase, in the present study this was observed only when logistic regressions were done that split children with high appetites into a separate group. In this case, child acceptance of food did rise during convalescence, suggesting that program messages should communicate to mothers the need to feed their children more during the recovery period. The data here show that mothers discontinued their more active encouragement to eat when diarrhea stopped, thereby missing an opportunity for their children to consume important "catch-up" nutrients. Program messages, therefore, should focus on a child's increase in appetite during the convalescent period, and encourage mothers to actively feed their children both during *and* after diarrhea. In addition, specific recommendations on foods appropriate during diarrhea and convalescence could be made and less emphasis placed on vague feeding messages. These messages should be based on cultural beliefs regarding appropriate foods to offer during diarrhea.[12]

We have shown a clear link between the presence of fever and the rejection of food by children. The effect of fever on appetite is worrisome because of the number of infections where fever is commonly present, such as upper respiratory infections and malaria. Lopez de Romaña, *et al*,[24] found fever present in children during 13.7 percent of all illness days, and Brown, *et al*,[25] reported an approximate 30 percent reduction in dietary intake among weanlings whose mothers reported the presence of fever, although Martorell, *et al*,[26] found no link between fever and growth rates. Clearly, more work is required to investigate the effect of fever on dietary intake and growth, and whether the administration of acetaminophen improves the acceptance of food by ill children (Gove, personal communication).

In conclusion, this study illustrates the importance of examining the behavioral aspects of nutritional intake. Simply measuring changes in caloric intake does not provide an explanation as to why the changes have occurred, so interventions may be based on faulty assumptions of behavior. In this case, mothers do not, as commonly assumed, withhold food from children with diarrhea. On the contrary, they actually increase their efforts to feed their children, as a response to the child's decreased acceptance of food. ☐

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The Epidemiology of Licit and Illicit Substance Use among High School Students in Greece

Anna Kokkevi, PhD, and Costas Stefanis, MD

ABSTRACT

Findings on self-reported adolescent licit and illicit substance use are presented based on a nationwide 1984 probability sample of 11,058 Greek adolescent students ages 14–18 years old. Regular smoking and use of alcohol in the 30 days prior to the survey were reported by 22.3 percent and 82.8 percent, respectively. Nearly one-third of the students (30.3 percent) reported lifetime use of psychotropic drugs without prescription, without ever having used any illicit drug. Illicit drug use is less common in Greece than in other European and North American countries with a lifetime prevalence rate of 6 percent. Regular use of tobacco, alcohol, and illicit drugs was more common in males, while unprescribed use of licit psychotropic drugs prevailed in females. Regular smoking, drinking, and illicit drug use were associated with urbanization but not socioeconomic status. Between 1984 and 1988 in an Athenian subsample of schools there was a 20 percent increase in illicit drug use among males. Use of psychotropic drugs without prescription decreased much more than use of tobacco and alcohol. *(Am J Public Health* 1991; 81:48–52)

Introduction

In Greece, police reports and data from other sources indicate that use of illicit drugs was spreading in the late 1970s to young persons of all socioeconomic classes. Front-page newspaper articles alarmed the public and campaigns against drugs were initiated by lay groups.

This situation led the authorities to acknowledge the need for assessing the situation accurately. The Department of Psychiatry of Athens University was assigned the task of studying the extent of the drug problem in Greece and submitting proposals for appropriate preventive action.

The present paper presents an overview of illicit substance use among adolescent students in Greece, as well as related aspects such as smoking, alcohol drinking, and the use of licit psychotropic drugs without medical prescription.

Trends in licit and illicit substance use within the four years (1984–88) are also presented for Athenian high school students.

Methods

Nationwide Study

The sample is a nationwide multistage stratified probability sample. Urban areas were divided into three strata (Athens, Salonika, and all other urban areas); a fourth stratum consisted of all semi-urban and rural areas. Stage 1 of the sampling procedure involved the selection of urban, semi-urban, and rural locations, stage 2 the selection of high schools, and stage 3 the selection of classrooms within each high school. The allocation of the sample among the strata was proportional to the size of their student populations. The selection of sampling units in the first and second stages was proportional to their size. In each selected school, one class of about 30 students was selected with equal probability from each grade, and all students in the selected classes comprised the final sample.

The total sample consisted of 11,058 students ages 14–18 years attending the last four grades of high school. The secondary school system (high school) in Greece consists of six years; the first three years (Gymnasium) are compulsory and the second (Lyceum, either general or technical-vocational education) non-compulsory. Approximately 80 percent of Gymnasium students continue their studies in the Lyceum. The majority of schools are run by the state; only 6 percent are private fee-paying. The age of students in the Gymnasium is 12–14 years and in the Lyceum 15–18 years, with negligible deviations from this range. All 96 selected schools, representing all types of high schools, participated in the survey.

Data collection took place from March to May 1984. Students answered an anonymous questionnaire administered in their classrooms during two normal class sessions. The questionnaire consisted of structured pre-coded items covering a wide range of areas in addition to the use of licit and illicit substances: sociodemographic, school-related, physical and psychosocial health, life styles, future goals, personality and attitudes. Fa-

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From the Department of Psychiatry, Athens University Medical School. Address reprint requests to Anna Kokkevi, PhD, Assistant Professor, Department of Psychiatry, Athens University Medical School, Eginition Hospital, Athens 11528 Greece. Dr. Stefanis is Professor of Psychiatry at the University. This paper, submitted to the Journal March 27, 1989, was revised and accepted for publication February 26, 1990.

ther's occupation and educational level as reported by the student were used in combination to assign students to three socioeconomic classes. This classification was validated by additional indices such as mother's educational level, housing, and estimated family income.

The average rate of absentees on the day of the survey was 6.1 percent. The participation rate of students present was 100 percent.

Reliability and validity of the responses concerning drug use were checked in a pilot study of 656 students conducted four months prior to the main study within the same school year. Students answered the questionnaire twice, with a 20-day interval, and questionnaires were matched by numbered stickers distributed at the first administration of the questionnaire.

Test-retest reliability coefficients were measured on 560 cases (85.4 percent of the total sample) for whom matching of the questionnaires was possible. Pearson's correlation coefficient for responses in seven categories ranged as follows: between .80 and 1.00 for tobacco smoking, drunkenness and lifetime and 12 month use of cannabis, cocaine, hallucinogens, and for 12-month use of heroin and opiates other than heroin; between .60 and .79 for lifetime and 12-month use of alcohol, stimulants, tranquilizers, hypnotics and for lifetime use of heroin and of opiates other than heroin, and .46 for lifetime use of heroin. Kappa coefficients[1] ranged as follows: between .71 and .89 for tobacco use and for lifetime and 12-month use of cannabis; between .41 and .50 for lifetime alcohol use, drunkenness, lifetime and 12-month use of hypnotics, for lifetime use of cocaine and opiates other than heroin and for 12-month use of tranquilizers; between .28 and .39 for 12-month alcohol use, lifetime and 12-month use of amphetamines, and lifetime use of heroin and tranquilizers. Reliability coefficients were significant at the .0001 level for 19 out of 20 usage variables examined and at .001 for the remaining one. No differences in reliability coefficients were observed between questions on drug use and other topics.

Validity was checked empirically by internal and external criteria:

- Rates of substance use and abuse in schools a priori known to have high rates of students with behavior problems and drug taking were compared with the rates reported in the remaining schools. Rates of reported experience with licit and illicit drugs proved to be substantially higher in the former schools for tobacco smoking, frequency of marijuana use in lifetime, last 12 months and 30 days, any use of heroin in the last 12 months and 30 days, any lifetime use of hallucinogens, opiates other than heroin, and lifetime and last 12 months' use of cocaine.
- 1.5 percent of the returned questionnaires were excluded from further analysis at the initial phase of checking because of major inconsistencies or unanswered parts of the questionnaires. In the retained questionnaires, internal consistency of answers was found to be very high, ranging from 85.5–100 percent.
- Urine was collected for the purpose of laboratory checking of general health status and substance use from 384 students in the four Athenian schools which participated in the pilot study. Students were informed of the general purpose of urine collection, but no specific mention was made of the substance tracing tests. Since anonymity was safeguarded, this procedure was in accordance with the ethical rules for investigations in Greece at the time. Urine samples were matched to questionnaires by the anonymous method of numbered stickers used in the test-retest procedure. Biochemical analysis for the detection of opiates, cocaine, and cannabis was performed by radioimmunoassay in the Instituto Superiori di Sanita (Rome). Results were positive for five of the 360 students for whom matching to questionnaires was possible: four were positive for cannabis and one for opiates. These five students had all responded positively to the pertinent questions but a further nine students (total rate 3.9 percent) reported use in the questionnaire without testing positive. Since reported use in the questionnaire covered the last 30 days while the bioassay techniques employed ensured detectability for only a few days following intake, some discrepancy in rates might be expected.

Study of Trends in an Athenian Sample

The sample in the 1988 survey consisted of classes from a randomly selected sub-sample of 18 schools from the 38 Athens area schools that had been included in the nationwide study of 1984. A total of 1,961 students participated in the survey. The average absentee rate on the day of survey was 5 percent.

A condensed version of the original questionnaire was used, covering only four out of eight categories of licit psychotropic drugs (tranquilizers, hypnotics, codeine, amphetamines) and three out of five illicit drugs (cannabis, heroin, cocaine). The same procedures were followed as in 1984 including test-retest reliability checks (135 matched pairs of questionnaires from 180 students over an interval of three weeks) and comparisons of rates of use in high-risk schools against the rest. Reliability and validity results were similar to those obtained in 1984.

Results

Nationwide Study

Over 70 percent of students reported having experimented at least once with tobacco smoking and 22 percent claimed regular smoking in the past month (Table 1). Regular smoking was highly correlated with sex and age, with more than one-third of the population smoking regularly in the older age group. The prevalence was much higher in the technical-vocational schools (41.9 percent vs 18.4 percent in public and 16.1 percent in private), in urban areas (23.7 percent vs 15.1 percent in semi-urban and rural areas), and lower in the higher socioeconomic level (19.4 percent vs 21.7 percent in the lower and 23.0 percent in the middle level).

The prevalence of alcohol use was 94.8 percent for past year and 82.4 percent for past month; 15.4 percent of students reported frequent alcohol use (10 or more times) in the past month. Rates of frequent alcohol use were twice as high in males as in females and nearly twice as high in the oldest as in the youngest age group (Table 1). Higher rates of frequent alcohol use were reported in private and technical-vocational in comparison to public schools (20.5 percent, 22.1 percent and 13.5 percent, respectively). Residence and socioeconomic level were not important determinants of the frequency of alcohol use. Heavy alcohol use (five or more drinks on one occasion) within the last 30 days was reported by 36.9 percent of students (44.4 percent male, 29.4 percent female) and increased from 32.2 percent in the 13–14 age group to 44.6 percent at 17–18 years.

Use of illicit drugs at any time was reported by 6 percent of students, with a ratio of two to one between boys and girls (Table 2). Approximately half of the users reported use more than twice.

The lifetime use of prescription requiring licit psychotropic drugs, without a doctor's order, was much higher than the use of illicit drugs, with a sex ratio of three girls to two boys. Nearly one-third (30.3 percent) of the student population reported that they had used unprescribed

TABLE 1—Tobacco and Alcohol Use by Sex and Age among Adolescents in Greece

		Sex		Age (years)		
	Total	Male	Female	13–14	15–16	17–18
Approximate Respondents (N)*	(10715)	(5354)	(5359)	(2708)	(5264)	(2734)
Lifetime smoking	%	%	%	%	%	%
Never	28.7	27.0	30.4	43.9	27.9	15.1
Once or twice	16.0	15.0	17.0	18.6	16.2	13.0
Occasionally	28.7	26.8	30.6	25.7	30.1	29.0
Regularly in the past	4.3	5.2	3.3	3.3	4.3	5.1
Regularly now	22.3	26.0	18.6	8.5	21.4	37.8
Past month frequency of alcohol use						
0 occasions	17.6	14.8	20.4	23.2	16.9	13.4
1–9 "	66.9	65.0	69.0	65.1	68.6	66.2
10–19 "	9.5	12.3	6.7	7.1	9.6	12.1
20–more	5.9	7.8	3.9	4.6	4.9	8.3

*Totals may differ because of small amounts of missing data.

TABLE 2—Lifetime, Past Year and Past Month Prevalences of Illicit and Unprescribed Psychotropic Drug Use (pills)* by Sex and Age

	Approximate Respondents (No.)	Illicit Drugs	Pills, not Illicit	Non-users
Lifetime	%	%	%	
Total rate	(10.812)	6.0	30.3	63.7
Sex				
Male	(5.396)	7.7	24.8	67.5
Female	(5.416)	4.0	35.7	60.3
Age (years)				
13–14	(2.741)	3.2	26.6	70.2
15–16	(5.312)	5.0	31.6	63.5
17–18	(2.752)	10.2	31.2	58.6
Past Year				
Total rate	(10.812)	3.8	16.5	79.7
Sex				
Male	(5.396)	4.8	12.4	82.8
Female	(5.416)	2.8	20.5	76.7
Past Month				
Total rate	(10.812)	1.7	6.3	92.0
Sex				
Male	(5.396)	2.3	5.5	92.0
Female	(5.416)	1.2	7.1	91.7

*Students were assigned according to their reported use of drugs to two mutually exclusive categories: (a) "Pills"—use *without* doctor's orders of licit psychotropic drugs requiring prescription (amphetamines, anticholinergics, tranquilizers, hypnotics, barbiturates, antidepressants, at least once and analgesics or codeine-containing syrups more than five times) *but no use of illicit drugs;* and (b) use of illicit drugs (cannabis, hallucinogens, cocaine, opiates other than heroin, heroin). In the last category students might have also used licit psychotropic drugs without doctor's orders.

pain relievers or codeine-containing syrups more than five times or another licit psychotropic drug at least once, but had never used any illicit drug. If users of illicit drugs are included, 34.7 percent had used a licit psychotropic drug without a doctor's orders.

The frequency of use of unprescribed licit drugs and illicit drugs is shown in Table 3. These very high rates of unprescribed licit drugs might to some extent be due to the possibility that the students, despite instructions, had included in these categories over-the-counter drugs for pain and cough relief, not requiring prescription.

Use of both licit and illicit substances increased significantly with age. High rates of illicit drug use were observed in males of the older age group (14.5 percent) and high rates of licit drug use in females of the older age group (40.6 percent). Illicit drugs were used in higher rates in technical-vocational schools (9.9 percent) compared to private (6 percent) and public (5 percent) schools.

The rate of illicit drug use was related to city size, ranging from 6.4 percent in Athens to 4.4 percent in semi-urban and rural regions. Rates of illicit use increased with socioeconomic level, although differences were less marked. The only demographic variables influencing use of licit drugs were sex and age.

Licit and illicit substance use was found to be highly correlated: 15.5 percent of regular tobacco smokers had used illicit drugs compared to 2.5 percent of the rest of the population; 10.5 percent compared to 3 percent had used three or more types of licit psychotropic drugs without doctor's prescription. Among heavy drinkers, 24 percent had used illicit drugs, compared to 2.1 percent of the rest of the population; 17.4 percent compared to 3.4 percent had used three or more types of licit psychotropic drugs without doctor's prescription. Of those who had used three or more types of psychotropic drugs without prescription, 77.8 percent had also used illicit drugs, compared to 3.9 percent of the rest of the population. Moreover, 69 percent of those who had used illicit drugs other than heroin and 82.4 percent of those who had used heroin reported use of licit psychotropic drugs without a doctor's prescription ("pills").

Trends in Licit and Illicit Substance Use in the Athenian Sample

Comparisons of rates of substance use are based on the 18 Athens schools which participated in both 1984 and 1988.

Age was defined differently in the two surveys (year of birth in 1984, age last birthday in 1988), but this should have had only a small effect on comparisons.

A decrease was found in regular tobacco smoking for boys and a lesser decrease for girls (Table 4). The decrease in regular smoking was most striking in the two younger age groups.

Alcohol consumption also decreased between 1984 and 1988 in both sexes. The decrease in alcohol use is mainly due to a reduction in the number of occasional users (those who had drunk less than 10 times in the last 30 days). The proportion of frequent users in the oldest age group actually increased slightly.

Lifetime prevalence of psychotropic substance use on doctor's orders showed a substantial decrease for codeine (from 45.6 percent to 22.5 percent), whereas no

TABLE 3—Frequency of Unprescribed Licit and Illicit Drug Use for Specific Drugs

Drug Use	Life time			Last 12 Months			Last 30 days		
	Total %	Times 1–2	Times ≥3	Total %	Times 1–2	Times ≥3	Total %	Times 1–2	Times ≥3
Pain relievers	38.5	15.8	22.7	28.1	14.4	13.7	11.6	10.8	0.8
Codeine	24.9	13.5	11.4	12.3	9.2	3.1	4.1	3.4	0.7
Tranquilizers	10.0	6.2	3.8	6.4	4.0	2.4	2.6	1.7	0.9
Hypnotics	4.1	2.8	1.3	2.4	1.6	0.8	1.3	0.9	0.4
Amphetamines	4.5	2.9	1.6	2.8	1.8	1.0	1.4	0.9	0.5
Barbiturates	3.4	2.1	1.2	1.6	1.0	0.6	0.9	0.5	0.4
Antidepressants	2.5	1.7	0.9	1.6	1.1	0.5	0.6	0.4	0.2
Anticholinergics	5.2	2.7	2.5	2.9	1.6	1.3	1.3	0.6	0.7
Cannabis	3.9	2.3	1.6	2.7	1.6	1.1	1.2	0.7	0.5
Hallucinogens	1.1	0.7	0.4	0.7	0.4	0.3	0.3	0.2	0.1
Cocaine	1.6	0.9	0.7	1.1	0.6	0.5	0.6	0.4	0.2
Heroin	0.6	0.4	0.2	0.5	0.3	0.2	0.2	0.1	0.1
Opiates (other than heroin)	0.9	0.6	0.3	1.6	0.4	1.2	0.2	0.1	0.1

TABLE 4—Use of Tobacco, Alcohol, Illicit Drugs and Unprescribed Psychotropic Drugs (pills)[*] in 1984 and 1988 by Sex and Age

	Sex				Age (years)					
	Male		Female		13–14		15–16		17–18	
	1984	1988	1984	1988	1984	1988	1984	1988	1984	1988
Approximate Respondents (No)	(1065)	(1007)	(1060)	(922)	(571)	(265)	(1009)	(930)	(545)	(736)
	%	%	%	%	%	%	%	%	%	%
Regular tobacco smoking	23.9	20.3	21.8	20.7	8.2	2.6	22.1	13.0	39.6	36.3
Past month alcohol use										
No use	15.2	22.1	16.2	31.7	21.2	37.8	15.2	28.0	10.9	21.0
Occasional use (1–9 times)	63.4	57.8	72.7	58.5	66.4	56.7	69.2	59.2	67.5	57.5
Frequent use (10–19 times)	13.8	12.3	7.3	7.1	7.8	4.1	10.4	9.4	13.7	12.4
Very frequent use (20+ times)	7.6	7.8	3.8	2.7	4.6	1.4	5.2	3.3	7.9	9.2
Lifetime drug use										
Illicit drugs	6.5	8.1	3.8	3.9	1.7	0	4.7	4.3	9.7	10.9
Pills but not illicit	23.3	12.8	37.3	20.0	28.5	13.6	35.7	19.9	37.3	21.6
Non-users	70.1	79.1	58.8	76.1	69.8	86.4	59.7	75.8	53.0	67.5

[*] Students were assigned according to their reported use of drugs to two mutually exclusive categories: (a) "Pills"—use *without* doctor's orders of licit psychotropic drugs requiring prescription (amphetamines, tranquilizers, hypnotics, codeine-containing syrups) *but no use of illicit drugs;* and (b) use of illicit drugs (cannabis, heroin, cocaine). In the last category students might have also used licit psychotropic drugs without doctor's orders.

significant change was found for tranquilizers, hypnotics, and amphetamines.

Lifetime prevalence of licit psychotropic drug use without doctor's orders ("pills") was markedly reduced in 1988, to almost half of the rate observed in 1984. On the other hand, lifetime prevalence of use of any illicit drug showed an increase of 20 percent for boys (from 6.5 percent in 1984 to 8.1 percent in 1988), while current use (past month) rates of illicit drug use almost doubled in both sexes. More specifically, lifetime rates of use increased from 6.1 percent to 7.5 percent for cannabis, from 0.9 percent to 1.7 percent for heroin, and from 1.3 percent to 2.3 percent for cocaine.

Discussion

The nationwide survey of Greek high school students shows a rate of illicit drug use much lower than the rates of 10 to 23 percent reported during the first half of the 1980s in similar studies in other European countries[2–6,*] and the rates of over 40 percent in the United States and Canada.[7–9]

Conversely, the overall use of licit substances such as tobacco and alcohol is higher in Greece than in the USA, Canada and northern European countries.[7,8,2,6,*] However, regular drinking is lower than in France,[10] and heavy drinking even lower than in the USA and Canada.[7,8]

The use of licit psychotropic drugs without a doctor's orders is much more common among the Greek students than the use of illicit drugs. In comparison to the US, the unprescribed use of opiates is more frequent in Greece, the use of tranquilizers is at the same level, and amphetamine and barbiturate use is less common.[7] Similar differences for the last three substances are found between Greece and Canada.[8]

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[*] Hibbel B: Trends in drinking and drug habits in Swedish youth. Paper presented at the 27th International Institute on the Prevention and Treatment of Alcoholism, Vienna, 1981.

The spread of licit and illicit substances among students in Greece is, as elsewhere, influenced by legal and societal factors. Greece is a producer of alcohol and tobacco and thus the use of these substances is widespread. The traditional pattern of alcohol use in Greece is cultural and recreational, with consumption during meals and festivities where drinking by younger members of the family is not generally discouraged. Although the availability of illicit drugs is rather high due to the fact that Greece is a pathway in the illicit trade between Asian countries and the West, the use of cannabis and other illicit drugs is considered as criminal behavior and subject to severe penalties. On the other hand strict regulations for prescribing and purchasing of psychotropics have been strongly implemented only recently. Access to psychotropic drugs was therefore rather easy at the time of the nationwide survey.

Trends in the last four years in the Athens area suggest a possible impact of policy and of social attitudes on the use of psychotropic medicines but little change in tobacco and alcohol use. Campaigns against unwarranted prescription of licit drugs and the unprescribed use of medicines, as well as the stricter implementation of laws regulating their provision, seem to have had a substantial effect on consumption of licit psychotropic drugs without prescription. The significant drop in rates of prescribed codeine use between 1984 and 1988 may also reflect increased awareness of medical professionals of the addictive aspects of the substance. Although there was some tendency toward the decline of regular tobacco and alcohol use among younger students and that of occasional alcohol use in all age groups, rates of more frequent use of alcohol were unchanged in the older age group. More frequent users may be more resistant to sociocultural pressure for change.

Greece does not seem to have escaped the first stage of the course of the epidemic of illicit drug use, before the leveling off or even the decrease of the phenomenon, as observed in other European and North American countries within the last decade. There has been a rapid expansion across barriers of age and social class although illicit drug use is still predominantly associated with males, as was also seen in a nationwide survey of the general population.[11] Along with the increased use of heroin and cannabis, the recent surge of cocaine use indicates that users in Greece are following closely the preferences and patterns of illicit drug use in western Europe and North America. ☐

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AIDS-Related Knowledge, Sexual Behavior, and Condom Use among Men and Women in Kinshasa, Zaire

Jane T. Bertrand, PhD, Bakutuvwidi Makani, MS, Susan E. Hassig, DrPH, Kinavwidi Lewu Niwembo, Balowa Djunghu, Mbadu Muanda, L2, and Chirwisa Chirhamolekwa, L2

ABSTRACT

This study was conducted in 1988 among a random sample of 6,625 men and women of reproductive age in all 24 administrative zones of Kinshasa, the capital city of Zaire, to determine existing levels of knowledge regarding AIDS (acquired immunodeficiency syndrome), sexual behavior, knowledge and use of condoms in marital and extramarital relations; perceived risk of AIDS, and attitudes toward testing for the human immunodeficiency virus (HIV). Awareness of AIDS is almost universal, and the vast majority know the four main modes of transmission. Almost half believed in transmission by mosquitoes and in a vaccine or cure for AIDS. The majority of male respondents knew of condoms, but negative attitudes toward condom use are widespread, and few respondents perceived them to play a central role in combatting AIDS. (*Am J Public Health* 1991; 81:53–58)

Acquired immunodeficiency syndrome (AIDS) constitutes a major public health problem in Zaire. It is estimated that 6–8 percent of the adult population of Kinshasa, the capital city of Zaire, are seropositive, based on serosurveys of selected populations.[1] Six percent of babies delivered in the two major hospitals in Kinshasa in 1987 were born to seropositive mothers.[2] In 1988, 50 percent of the admissions to the internal medicine ward of the major public hospital, Mama Yemo, were HIV (human immunodeficiency virus) seropositive.[3] It is estimated that 80–90 percent of AIDS in Kinshasa is transmitted through sexual contact.[1] Since homosexuality is rare in Kinshasa, heterosexual relations constitute the major means of AIDS transmission among this population.

Until a vaccine or cure is found, public information and education remain an important weapon against AIDS. Since 1987, AIDS messages have been broadcast frequently on radio and television. Luambo, a popular singer in Zaire, further publicized the dangers of AIDS through a widely diffused song.

The current study was conducted city-wide in 1988 among the adult population to measure awareness of AIDS, its modes of transmission and means of prevention; coital frequency and extramarital sexual relations; knowledge and use of condoms; perceived risk of getting AIDS; and attitudes toward HIV testing. This module on AIDS formed part of a larger survey on contraceptive prevalence in Kinshasa, conducted by the Projet des Services des Naissances Désirables, the national family planning services project.

Methods

The projected sample for this survey consisted of 1,500 men (ages 20–59 years) and 1,500 women (ages 15–49 years), randomly selected from the 24 administrative zones of Kinshasa. The age criterion differed by sex, in an effort to obtain the population at greatest risk of AIDS. Four zones were intentionally oversampled for other research purposes (a field study to test the impact of an AIDS prevention program, not reported herein), bringing the projected sample size to 2,500 per sex.

Two-staged sampling was involved: first, 72 of 251 *quartiers* (the largest administrative unit within an administrative zone) were randomly selected and subsequently mapped; second, *parcelles* (a compound containing one or more households) were systematically selected. Men and women were interviewed in alternate *parcelles*; thus, the male/female data constitute independent samples, not husbands and wives.

Address reprint requests to Jane T. Bertrand, PhD, Associate Professor, Department of Applied Health Sciences, Tulane University, School of Public Health and Tropical Medicine, 1501 Canal Street, New Orleans, LA 70112. About the co-authors: Bakutuvwidi is Consultant, Centre d'Informations Techniques et de Recherches pour le Développement, Kinshasa, Zaire; Dr. Hassig is Research Assistant Professor, Biostatistics and Epidemiology, Tulane SPHTM, New Orleans; Kinavwidi is Statistician, Institut National de la Statistique, Kinshasa; Balowa is Head of Computer Services, Mbadu is Head of Operations Research, and Chirwisa is Director, all three with Projet des Services des Naissances Désirables (the Family Planning Services Project), Kinshasa. This paper, submitted to the Journal June 19, 1989, was revised and accepted for publication March 21, 1990.

TABLE 1—Sociodemographic Characteristics of the Study Population*

	Men N = 3140	Women N = 3485
Age (years)	%	%
15–19	—	23.0
20–24	23.5	23.2
25–29	36.2	20.0
30–34	12.8	14.6
35–39	9.8	10.2
40–44	9.6	5.6
45 and over	8.1	3.2
Mean age (years)	33.4	26.6
Highest educational level attained		
None or primary	19.1	36.4
Secondary	61.0	60.5
Beyond secondary	19.5	2.7
Marital status		
Married (1 wife)	53.1	44.1
Married (2+ wives)	4.3	8.3
Consensual union	2.0	9.5
Separated, divorced, or widowed	3.6	8.5
Single, never married	37.0	29.5
Household amenities		
Has electricity	33.6	35.3
Has running water in house	9.2	7.8
Has toilet inside house	11.5	8.6
Has a working radio	51.8	49.3
Has a car or truck	9.4	7.0

*The "N" given in this and all subsequent tables is based on the unweighted sample (i.e. actual number of people who answered the question); the percentages are based on the weighted sample.

Because the average number of eligible respondents per *parcelle* was higher than expected, the total number of respondents (6,625) exceeded the projected sample. In the tables below, the results have been weighted to adjust for oversampling. The n in each case is based on the unweighted sample (actual number of people who were asked the question), whereas the percentages are based on the weighted sample.

Data collection for this survey took place from January to September 1988. All interviewers underwent a 20-day training course on sampling and data collection procedures prior to the fieldwork. All interviews were conducted in Lingala, the local language, by an interviewer of the same sex, using a pretested, structured questionnaire. Each interview lasted an average of 30-40 minutes.

All questionnaires were verified for completeness and logical consistency; where necessary, interviewers were sent back to the field to verify or complete responses. SPSS/PC + and BMDP/PC were used for the analysis.

Results

The total sample included 3,140 men and 3,485 women (Table 1). Over half of all respondents had at least some level of secondary education; however, significantly more men (20 percent) than women (3 percent) had gone beyond secondary school. The majority of men (59 percent) and women (62 percent) were married or living in consensual union.

Knowledge of AIDS

Awareness of AIDS is almost universal in Kinshasa (99 percent of men and 96 percent of women). Details of specific AIDS knowledge items are presented in Table 2. Of note, only half the study population knew that mosquitoes do not transmit AIDS and only 70 percent had correct information regarding casual contact. Ninety-four percent of men and 88 percent of women believed that AIDS could be prevented through behavioral changes.

Thirty-six percent of men and 45 percent of women at the time of this survey believed that there was a vaccine against HIV. Moreover, 40 percent of men and 21 percent of women believed that a person with AIDS can be cured. An additional 5–12 percent were unsure regarding both of these points.

Seventeen percent of men and 31 percent of women knew someone with AIDS at the time of the interview. Half of all respondents knew at least one person who had died of AIDS.

All respondents who had heard of AIDS were asked whether they had seen or heard something about AIDS via different communication channels. More adults in Kinshasa have been reached via radio than any other medium: 96 percent of men and 83 percent of women. Almost identical percentages of each group reported having heard the Luambo song on AIDS. A slightly lower percentage (86 percent of men, 70 percent of women) had seen something about AIDS on television. The percentages who had seen brochures, newspaper articles, or posters on AIDS were lower, ranging from 38–60 percent for men and 22–34 percent for women.

The vast majority of respondents, 95 percent of men and 89 percent of women, wanted to learn more about AIDS.

Multivariate Predictors of Knowledge of AIDS

In a multiple regression of the summed knowledge score by demographic and descriptive variables, males and females both had similar results (Table 3). Higher knowledge scores in both groups were associated with higher socioeconomic status (SES), higher educational attainment, length of time and number of messages they had heard about AIDS, personally knowing someone who has AIDS or had died of AIDS, and professing to a non-Christian religion. Being raised in a rural environment was negatively associated with knowledge for women, but not for men. Being older was negatively associated with knowledge for men but not for women. (In both males and females, age was negatively correlated with level of education.) Despite the number of significantly associated variables, the adjusted r square for both the male and female models was only 0.18. Thus, it is apparent that the variability seen in the knowledge scores in this population cannot be truly defined by the available variables.

Sexual Behavior, Knowledge, and Use of Condoms in Marriage

The median age at first sex (among those who had initiated sexual relations) was 17 years for men and 16 years for women. One percent of men and 14 percent women had not had sex, while 9 percent men and 4 percent women did not respond to the question.

All married respondents (N = 4,008) were asked whether they had had sexual

TABLE 2—AIDS Knowledge Items and Perceived Risk of HIV Infection among Those Who Have Heard of AIDS (percent with correct response)

	Men N = 3099	Women N = 3351
The AIDS virus is transmitted by:		
• eating unclean food	82.6	76.6
• transfusion with infected blood*	96.8	92.7
• mosquitoes	50.5	55.3
• sexual relations with an infected person*	98.7	96.4
• living in the same house with AIDS patient	66.5	56.6
• injection with unsterile needle*	95.7	91.3
• wearing used clothing	71.0	63.7
• working beside an AIDS victim	71.3	62.4
• (infected) mother to fetus*	94.0	89.2
The following are symptoms of AIDS:		
• frequent vomiting	63.5	63.5
• prolonged diarrhea*	81.6	81.7
• heavy bleeding for women	63.3	63.9
• recurrent fever*	45.2	43.2
• extreme weight loss*	78.2	68.7
• loss of hearing	75.2	73.6
• pain in the joints	57.1	56.1
• general, prolonged fatigue*	43.4	28.8
Transmission beliefs:		
• person can be infected without knowing*	87.9	77.5
• person can be infected without feeling sick*	84.5	73.1
• person who feels healthy could transmit the AIDS virus to another person*	84.2	72.1
• person with multiple partners has more risk than person with one partner*	78.9	72.6
• it is advisable for a woman infected with AIDS virus to get pregnant	97.4	95.7
• person infected with the AIDS virus will eventually get the disease*	76.0	71.2
There is a vaccine againt AIDS	58.6	43.1
AIDS can be cured	52.0	68.0
AIDS can be prevented*	93.7	88.0
AIDS can be prevented by:		
• taking antibiotics before sex	81.5	65.4
• using condoms for extramarital sex*	70.1	41.6
• going to a traditional healer	90.3	79.1
• avoiding sex with multiple partners*	93.7	85.4
Perceived risk of contacting AIDS:		
• Very high	10.6	10.2
• Moderate	6.8	5.9
• Low	21.0	9.3
• None	46.8	62.9
• Don't know	13.8	11.3

The asterisks above indicate which items are correct as stated.

relations in the previous month; 66 percent of 1,885 men and 60 percent of 2,123 women reported having had relations during this period. Abstinence was associated with the postpartum period or having an absent partner. Among those married respondents who had had relations, the mean number of days of sex during the previous week was 1.3.

Over 90 percent of married and single men, 77 percent of married women, and 61 percent of single women had heard of condoms. Despite the high recognition level, relatively few of the respondents in this survey had ever used condoms, with 12 percent of married men reporting ever using them with their spouse.

Eight percent of married men and 7 percent of married women reported current condom use with their spouse, although less than 2 percent of either sex reportedly used them "always" or "most of the time." The predominant reason for using condoms with one's spouse was to prevent pregnancy; the second, to prevent sexually transmitted diseases (STDs) in general, and the third to prevent AIDS. Mention of pregnancy prevention was higher among married than single respondents (reporting on condom use with their current partner), whereas singles cited prevention of other STDs and AIDS more often than married respondents.

Extramarital Relations and Use of Condoms

Twenty-three percent of married men and 1 percent of married women reported having had an extramarital partner or partners in the past six months; 12 percent of men and 1 percent of women reported such relations in the four weeks preceding the interview. Of men having had extramarital activity in the past six months, half reported having only one such partner (Table 4). Among the relatively few women who admitted to extramarital sexual activity, almost all claimed that it was with a single partner.

Factors associated with extramarital activity in men included younger age (<35 years), moderately higher education and SES levels, increased belief in the existence of a vaccine against HIV, and residence in a major urban area since adolescence. Men acknowledging extramarital activity also had slightly more misconceptions regarding false modes of HIV transmission, had heard more messages regarding AIDS than those denying extramarital activity, were more likely to cite condoms as a prevention method for AIDS, to perceive themselves to be at "very high" or "high" risk of AIDS, and to have known someone who had died of AIDS.

In women, the factors linked with such activity were more difficult to define, given the small numbers of women who acknowledged the behavior, but a few basic factors were similar to the men: younger age and urban residence since adolescence were associated with extramarital activity.

While the numbers are admittedly small, condom use was higher with extramarital partners (24 percent among men, 12 percent among women) than with one's spouse. In addition, the frequency of use was more consistent with extramarital partners than with spouses; half of the male users did so "always" or "most of the time." Among unmarried men, 27 percent reported having used condoms with their current partner, with a frequency of use similar to married men in their extramarital relations.

Attitudes toward Condoms

Over 52 percent of all male respondents believed that condoms tear easily during sexual relations, that they can stay in the vagina after sex, and that they diminish sexual pleasure. An additional 12–28 percent had no opinion on these questions. Single men were slightly more favorable toward condoms than married men.

TABLE 3—Multiple Regression of Summed Knowledge Score by Demographic Variables*

	Male		Female	
	Coeff	SE of Coeff	Coeff	SE of Coeff
Socioeconomic status	0.4488	0.0706	0.3709	0.0798
Age	−0.0295	0.0116	−0.0025	0.0136
Non-Christian religion	−0.2165	0.1008	−0.4542	0.1157
Multiple partners or extramarital partners	−0.0503	0.0680	0.6069	0.5444
Time source heard of AIDS/HIV	0.3650	0.0608	0.2405	0.0636
Number of information channels	0.4384	0.1027	0.8820	0.1240
Know someone with/died of AIDS	0.5844	0.1859	0.6707	0.2181
Rural residence	0.0384	0.1062	−0.3273	0.1245
Level of education	0.7765	0.1284	1.0612	0.2135
Marital status	0.2859	0.2275	0.1092	0.2193
	Adjusted $R^2 = 0.18$		Adjusted $R^2 = 0.18$	

*The knowledge score is based on one point for each correct answer to the items listed in Table 2 (except for "perceived risk").

TABLE 4—Frequency of Extramarital Relations and Condom Use among Married Respondents Who Have Heard about AIDS

	Men N = 1885	Women N = 2123
Number of extramarital partners in past 6 months		
0	77.4	98.6
1	11.7	1.3
2	5.6	0.1
3	2.9	-
4 or more	2.3	-
Percentage reporting extramarital relations in past four weeks	12.1	1.2
Among those with extramarital partners in past 6 months	(N = 401)	(N = 22)
% who use condoms with extramarital partners	24.0	11.8
Among these condom users the frequency of use	(N = 90)	(N = 4) (N too small to report %)
Always	35.4	
Most of the time	14.5	
Sometimes	27.4	
Never	22.7	

Among women, 30–50 percent believed that condoms have the above-mentioned disadvantages. However, well over 30 percent of both married and single women answered "don't know" to these questions, which could reflect their ignorance of the method, lack of experience with condoms, or embarrassment over the questions. Married women were slightly more knowledgeable or favorable than single women (data not shown).

Factors associated with condom use and with relatively positive attitudes toward condoms for men were younger age (<35), having had extramarital activity or multiple partners in the last six months, and not having a steady partner. Men with positive attitudes toward condoms were more likely to have used condoms and to cite condoms as a means of HIV prevention. Neither SES, education, religion, nor urban residence showed any association with condom attitudes (data not shown).

Perceived Risk of Getting AIDS

Table 2 shows the distribution of perceived risk in the male and female study populations. What is most striking is that 47 percent of men and 63 percent of women believed that they were at no risk of getting AIDS. For men, factors associated with an increased perception of risk include younger age, higher education and SES levels, knowing someone who died of AIDS, having less accurate knowledge regarding false modes of HIV transmission, having heard more messages regarding AIDS, being either Catholic or Protestant, and having had multiple sexual partners or engaged in extramarital activity in the last six months.

An increased perceived risk in women was associated with higher education and SES levels, engaging in extramarital activity or having multiple sexual partners, having heard more messages regarding AIDS and knowing of AIDS for a longer period of time, receiving their AIDS information from informal sources such as friends and neighbors, and knowing someone who died of AIDS. Women with high perceived risk also had less knowledge regarding common means of AIDS prevention and were slightly more likely to recognize that there is no available vaccine. Surprisingly, women who had a stable or regular partner also perceived themselves to be at increased risk versus those women without such a partner. This may be a reflection of female awareness of or concern over male infidelity.

The association of perceived high risk with misconceptions regarding modes of AIDS transmission and means of prevention suggests that at least a part of this perception may not be based on fact. However, considerable denial regarding AIDS risks is also playing a role in this population. Although having multiple sexual partners is a well-recognized risk behavior for AIDS, 74 percent of men actually engaging in extramarital relationships perceive themselves to be at little or no risk of AIDS.

Attitudes toward Testing for the AIDS Virus

Thirty-nine percent of men and 16 percent of the women had heard of the term "seropositive"; two-thirds of those who claimed to know the word defined it correctly. The majority of respondents, 78 percent of men and 57 percent of women, reported that they would accept being tested for HIV. Over 90 percent of respondents would want to be informed if either they or their spouse or child were diagnosed with AIDS or HIV infection.

TABLE 5—Perceived Risk of AIDS and Correlates of High Perceived Risk

	Men N = 3099	Women N = 3351
Perceived risk of contacting AIDS:	(%)	(%)
High	10.6	10.2
Moderate	6.8	5.9
Low	21.0	9.3
None	46.8	62.9
Don't know	13.8	11.3
Selected correlates of increased risk of AIDS (% with high or moderate response)		
Know person who has died of AIDS		
Yes	22.6	23.5
No	17.9	10.9
(p-value)	(0.03)	(<0.00000)
Multiple sexual partners in last 6 months		
Yes	26.3	32.6
No	16.4	19.9
(p-value)	(0.00001)	(0.05)
Number of AIDS messages heard		
1–3	10.7	16.5
4 or more	22.3	24.5
(p-value)	(<0.00000)	(0.0002)
In relationship with a stable partner		
Yes	19.7	21.2
No	21.0	14.4
(p-value)	(NS)	(0.0005)

Discussion

The results of this study provide the first available data on levels of knowledge and attitudes with regard to AIDS among a probabilistic sample of the adult population of Kinshasa. While awareness and knowledge may have increased since the study ended in September 1988, the results nonetheless constitute a useful basis for the planning and evaluation of future information-education-communication (IEC) activities for AIDS.

Awareness regarding transmission of HIV via sexual contact is very high, but misconceptions persist over transmission by casual contact and mosquitoes. The fact that only a relatively low percentage could identify symptoms of AIDS may be explained by the fact that health officials have not stressed these aspects in their educational messages.

The misconceptions among a significant portion of the population regarding a vaccine and cure for AIDS are reason for concern, since they may create a sense of false security. The belief regarding the vaccine may result from periodic news stories about vaccine research ongoing in Zaire. The belief that AIDS can be cured may also be attributable to the widespread publicity about a new drug "MM1," which was presented to the Zairian public in November 1987 as a treatment for AIDS. Two recent surveys in different populations have shown a drop over the past year in the percentage that believe in a vaccine and/or cure.[4,*]

Comparisons between the Kinshasa population and a May-June 1988 nationwide survey of the adult population in the United States[5] reveal that both populations were highly knowledgeable on modes of transmission and reported a high level of exposure to AIDS messages via the mass media; in Zaire the most far-reaching channel was radio, whereas in the US it was television.

However, the two populations differed markedly in several other respects: belief in a vaccine or cure for AIDS was far more common in Zaire, while belief that condoms were effective prevention measures was less common in Zaire. These differences also underscore the areas in which campaigns in Zaire will need to be more explicit.

This study is the first to provide data on two key variables related to HIV transmission—extramarital relations and condom use—from a large, probablistic sample of both men and women in Zaire. The responses to the questions on sexual behavior should be interpreted with some caution, however, as such items are highly sensitive.[6] Based upon anecdotal evidence from Kinshasa residents, it was initially believed that the current data underestimated both coital frequency (median number of days of sex per week: 1.3 for married respondents) and extra-marital relations (23 percent of married men and 1 percent of married women in the past six months).

However, our findings are similar to results obtained from a study of over 12,000 workers and their spouses in two businesses in Kinshasa.[7] Among, 4,487 married couples, the mean number of days during the past year in which the couples had sex was 84; this yields a mean of 1.6 days per week. Regarding extramarital sex, 29 percent of married men reported at least one extramarital partner in the past 12 months. This percentage is slightly higher than that found in our study but covers a 12-month rather than six-month period. The question of extramarital relations was posed to females at only the smaller of the two study sites and yielded a prevalence similar to ours (<2 percent).

By contrast, in a study of health personnel (ranging from doctors to traditional practitioners) from all regions of Zaire except Kinshasa, the percentage of men reporting a history of extramarital relations in the past was 58 percent in 1987, dropping to 43 percent in 1988; among women, these percentages were 29 percent in 1987, decreasing to 18 percent in 1988.[8] These percentages are substantially higher than those found for the Kinshasa population in general. Possible explanations for these differences could be that the time period for the activity was not clearly specified, the health workers had greater economic means with which to engage in such activities, or they felt more comfortable with the circumstances in which the information was obtained.

We believe that the data from the current study underestimate the extent of extramarital relations, especially among women. In a subsequent study of sexual behavior in Kinshasa based in part on focus groups, both male and female participants were asked whether married men and married women would be likely to answer truthfully to questions regarding extramarital relations on a survey. The general consensus was that it depended on the interviewer's skill at putting the respondent at ease. However, there was strong consensus that a married woman

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[*] Bertrand J, Bakutuvwidi M, Balowa D, Kinavwidi N: Knowledge of AIDS, sexual behavior and condom use in ten sites in Zaire. American Public Health Association annual meeting, Chicago, October 1989.

could not risk revealing anything about extramarital affairs, lest this get back to her husband."**

The findings with regard to knowledge and use of condoms among married women concur with previous studies in Zaire.[9-11] In addition, the data from the current study confirm the widespread supposition that condoms are used much more frequently for extramarital sex than for relations with one's spouse. Despite the fact that relatively few men report having used condoms, the majority claim that they tear, stay inside the women's vagina, and decrease sexual pleasure. These negative attitudes must be based on conjecture rather than experience. Unsubstantiated rumors have apparently discouraged the use of condoms to such a degree that the Zairian public does not perceive their importance in the fight against AIDS. The percentage of married men and married women reporting current condom use yielded very similar results from two independent samples. This refutes the hypothesis that women tend to underreport condom use (at least for relations with one's spouse), an issue which has been raised in connection with female-only contraceptive prevalence surveys.

Regarding perceived risk, most of the individuals engaging in extramarital and multiple partner sexual activity did not acknowledge any perception of risk, even though the majority of these same people intellectually "know" that sexual activity with multiple partners provides an elevated risk of HIV exposure. In addition, many persons with high perceived risk believed in the possibility of mosquito-born and/or casual transmission. An unfounded perception of risk, such as this, may be counterproductive in the course of HIV infection and AIDS. If people persist in fearing casual contact with infected individuals, the subsequent ostracism of such persons will place an unmanageable burden upon limited governmental resources. It may also limit the population's response to efforts to promote behavioral change.

To conclude, the findings from this survey have several important implications for ongoing and future efforts to prevent the spread of AIDS in Kinshasa:

- Greater efforts are needed to combat the still prevalent misconceptions that AIDS can be transmitted by casual contact and by mosquitoes.
- IEC efforts should reinforce the idea that the two main means of protection from sexually transmitted HIV infection or AIDS are total mutual fidelity within monogamous couples and where that is not considered possible, use of condoms.
- Condom promotion should include messages directed at changing the negative attitudes toward condoms reflected in this study.
- Information campaigns should emphasize the fact that there is no effective vaccine against AIDS and that the disease is fatal.
- Education campaigns should be directed at teens as well as adults; by the age of 17, half of males and females have engaged in sexual relations. ☐

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The Course of the HIV Epidemic among Intravenous Drug Users in Amsterdam, The Netherlands

Harry J.A. van Haastrecht, MS, Johanna A.R. van den Hoek, MD, PhD, Christiane Bardoux, MD, Anne Leentvaar-Kuypers, MD, and Roel A. Coutinho, MD, PhD

ABSTRACT

To determine if behavioral changes in intravenous drug users in Amsterdam have retarded the HIV (human immunodeficiency virus) epidemic in this group in recent years, we report that: HIV-antibody seroprevalence in annual samples of injectors has been constant over the years 1986–89; HIV-antibody incidence in a cohort of injectors appears to have decreased from 1986 to 1987 and stabilized after that until 1989; acute hepatitis B incidence in all drug users in Amsterdam declined rapidly between 1985–89. It is concluded that changes in drug use behavior so far appear to have resulted in a stabilization of the epidemic among injectors, at a level with a still disturbingly high incidence rate of 5–6 per 100 person-years. (*Am J Public Health* 1991; 81:59–62)

Introduction

Intravenous drug use was the only risk factor in 75 (7 percent) of the 1,074 AIDS (acquired immunodeficiency syndrome) cases diagnosed and reported in The Netherlands by December 31, 1989. Forty-five of these cases were residents of Amsterdam (P. Bindels, personal communication). Although the absolute number of AIDS cases among intravenous drug users in this city of approximately 700,000 inhabitants is still relatively small, the rate of the recent increase is high: from one new case in 1985, three in 1986, eight in 1987, and 15 in 1988 to 18 new cases in 1989. In an effort to stop or slow down the spread of HIV among drug users, a number of national and local policies have been developed and applied. The informational programs directed to drug users, on the dangers of AIDS and on ways of preventing infection (since 1986), were supplemented by supplying condoms to addicted prostitutes. In 1984, a needle and syringe exchange program was implemented,[1,2] the original purpose of which was to prevent the spread of hepatitis B. All of these policies aim at voluntary behavioral change, by reducing risk behavior and thereby limiting the occurrence of new HIV infections. Evidence based on self-reporting suggests that behavioral change toward safer drug using practices did indeed occur in Amsterdam in recent years. A study previously published in this Journal,[3] showed that in a cohort of intravenous drug users a marked increase in making exclusive use of the Amsterdam needle and syringe exchange program occurred during the period December 1985 through July 1988 with a concurrent marked reduction in sharing used needles and syringes. Self-reported changes in behavior have their limitations and more objective measures of risk reduction are needed to corroborate these findings. The aim of this study therefore is to describe several of these more objective indicators for the development of the HIV epidemic among drug users in Amsterdam during the past few years.

Three indicators will be dealt with: 1) HIV seroprevalence among intravenous drug users at intake into a study from 1986 through 1989; 2) incidence of HIV infections in a cohort of intravenous drug users from 1986 through 1989; and 3) incidence of acute hepatitis B infection among drug users in the city of Amsterdam from 1983 through 1989. In drug users, hepatitis B virus (HBV), like HIV, is transmitted mainly through the sharing of needles and syringes. Acute hepatitis B infection was therefore chosen as a surrogate marker for HIV infection. In homosexual men, where both HBV and HIV also share the same main transmission mode (anogenital intercourse), a close correspondence between hepatitis and HIV incidence was shown.[4]

Methods

Study Participants

Participants were recruited at six methadone maintenance outposts and at a weekly sexually transmitted diseases

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From the Municipal Health Service, Amsterdam, The Netherlands. Address reprint requests to H.J.A. van Haastrecht, Municipal Health Service, Department of Public Health and Environment, PO Box 20244, 1000 HE Amsterdam, The Netherlands. This paper, submitted to the Journal March 21, 1990, was revised and accepted for publication September 5, 1990.

TABLE 1—HIV-antibody Prevalence and Presence of Independent Risk Factors in Annual Samples of IVDU in Amsterdam

Independent Risk Factors*	1986 (N = 220)	1987 (N = 178)	1988 (N = 133)	1989 (N = 91)
% first injection > 5 years ago	75.0	73.4	66.2	71.4
% last injection in past week	61.2	57.6	58.6	57.1
% German nationality	20.0	18.5	16.5	17.6
% residing in Amsterdam > 2 years	72.3	75.8	72.2	78.0
% smoking of heroin not daily	84.5	78.0	83.5	76.9
% HIV-ab positive†	33.2	30.3	27.8	31.9
% adjusted **HIV-ab positive	33.2	31.0	31.2	34.1

*Presented categories are associated with a positive HIV-ab test.

**Calculated, using regression coefficients for the different years from a logistic model where all listed risk factors were included, with 1986 as arbitrary year of reference.

†χ^2 test for trend p = 0.54

clinic for drug-using prostitutes. Both intravenous and non-intravenous drug users were invited to participate. After informed consent was obtained, a blood sample for HIV serology was taken and participants were interviewed using a standard questionnaire. Participants were asked to return for follow-up examinations every four months for repeat interviews and blood samples. There was no payment for their first visit; those who continued participation in the study were offered 25 Dutch guilders as an incentive for each follow-up visit. HIV antibody tests were performed using standard procedures described elsewhere.[5] The intake of participants started December 1985 and continues to the present.

Results reported here pertain to the period January 1986 to the end of 1989. Participants who reported at intake ever having used drugs intravenously were considered intravenous drug users (IVDUs) and were all included in the HIV-seroprevalence analysis. The HIV-incidence analysis was limited to those follow-up participants who had used drugs intravenously within the five years preceding intake or who (re-)started intravenous drug use in the course of follow-up. IVDUs with a positive HIV-antibody test result at entry were more inclined to take part in the follow-up than those who tested negative; 69 percent versus 57 percent, respectively, returned for at least one follow-up visit by March 1990. Risk factors for the presence of HIV-antibody that were more often present in those in follow-up than in others were: recently injecting, residing in Amsterdam for more than two years, and German nationality.

HIV-Seroprevalence and -Incidence

The development of the HIV-seroprevalence rate over time was examined by separating the intakes according to year of intake (1986–89). Since participants were self-selected, it is conceivable that IVDUs with different probabilities of being seropositive entered the study at different times. To control for such a selection bias, predetermined risk factors[5] that predict the probability of being seropositive and that were presumed not to reflect policy-induced behavior changes were multivariately controlled by logistic regression, using standard statistical software.[6] These risk factors were: date of first and date of last injection, nationality, length of stay in Amsterdam, and frequency of smoking of heroin in past six months. Annual incidence rates were calculated by determining the number of seroconversions per 100 person-years for those at risk. At risk were considered IVDUs who were HIV-antibody negative at entry, from their first visit until their last one or until seroconversion. Date of seroconversion was estimated as the midpoint of date of last seronegative visit and date of first seropositive visit.

Incidence of Acute Hepatitis B

All cases of acute hepatitis B in Amsterdam must be reported to the Municipal Health Service. On a routine basis, background data including drug use are collected in a personal interview by a public health nurse. This procedure remained unchanged throughout the study period. All records of acute hepatitis B cases from 1983 through 1989 were carefully screened by one of us (CB) for cases with a history of illicit drug use. Incidence was calculated by using estimates of the total number of hard drug users present per year as denominator; these estimates were calculated by the Municipal Health Service's drug department, based on a capture/recapture method quarterly[7] and have wide margins of uncertainty. We assumed there was a constant proportion of injectors among these drug users during the study period.

Results

HIV-Seroprevalence

A total of 622 subjects who had ever used drugs intravenously participated in the study between January 1986 and December 1989 (Table 1). In four consecutive years (1986–89), both the crude HIV seroprevalence rates and the adjusted rates (taking 1986 as year of reference) remained between 27.8 percent and 34.1 percent with no trend effect (Table 1).

HIV-Incidence

A total of 236 subjects who were seronegative at entry into the study and who reported intravenous drug use either within the five years preceding entry or in the course of follow-up, were analyzed to determine incidence. Twenty-seven seroconversions were recorded: five in 1986, six in 1987, eight in 1988, and eight in 1989.

The incidence rates per 100 person-years (95% confidence interval) were: 10.7 (1.3, 20.0) in 1986, 5.6 (1.1, 10.1) in 1987, 5.6 (1.7, 9.5) in 1988, and 5.5 (1.7, 9.4) in 1989 (Figure 1). The incidence rate difference between 1986 and 1987 was 5.0 (−5.3, 15.4).

Acute Hepatitis B Incidence

The incidence of acute hepatitis B (AHB) per 10,000 varied from a high of 48 in 1984, to a low of 8.6 in 1989 (Figure 2). Over the seven-year time span under study, there was a significant decline in AHB incidence (χ^2 for trend = 23.5, df = 1, $p = 10^{-6}$).

Discussion

These data suggest a decline in HIV-seroconversion from 1986 to 1987, followed by stabilization up to 1989. The confidence intervals, however, are broad, indicating that the early decline is not conclusive. In San Francisco, a similar decline was noticed in HIV seroconversion rate per 100 person-years in IVDUs: from seven in 1985 to 2.4 in 1987.[8] It is very difficult to evaluate properly the impact of behavioral change among IVDUs in Amsterdam, by comparing it with studies being done in other countries. Generally, because large segments of the total population of IVDUs are hard to reach, very little is known about the representative-

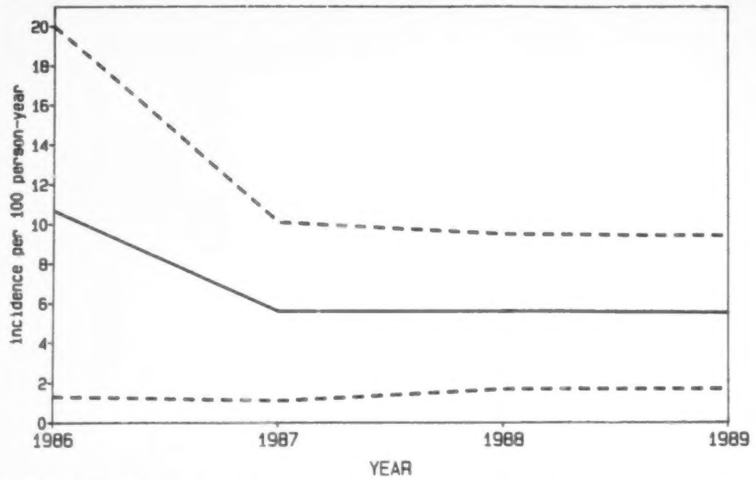

FIGURE 1—Incidence density and 95% confidence interval of HIV-antibody seroconversion in IVDU per year, 1986–89

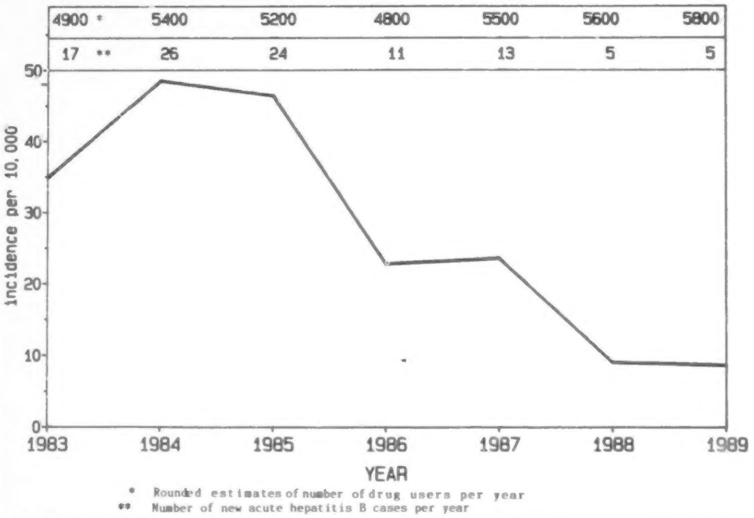

FIGURE 2—Estimated incidence of acute hepatitis B per 10,000 drug users in Amsterdam, The Netherlands, 1983–89

ness of IVDU study samples; ours is no exception. Participants were self-selected and follow-up was poor; only 57 percent of HIV-antibody negative IVDU at entry returned for at least one follow-up visit. Continued participation in the study may be of more interest to those IVDUs who have not modified their risk behavior and therefore may suspect a change in their serostatus since their last test. For this reason we may be overestimating HIV incidence in IVDUs in Amsterdam when generalizing from this group. On the other hand, we have evidence that behavioral change did occur especially in those drug users who took part in the follow-up study.[3]

We found rather stable HIV seroprevalences over the years, both for the crude prevalences and after attempting to correct for the influence of selection biases following from our sampling procedures. Stable HIV seroprevalences among IVDUs in recent years have been reported from many cities in North America and Western Europe.[8–16] The interpretation of this phenomenon, however, and especially measuring the impact of behavior change by monitoring seroprevalence, is very difficult since many factors that affect the general IVDU population and that are generally not well recorded—for example loss of seropositives for various reasons (mortality, hospitalization, moving, cessation of drug use, etc.) and entrance of new seronegative injectors—have to be taken into account. As appears from our data and as was reported before from Manhattan, New York City,[9] a stable seroprevalence is still compatible with a rather high incidence.

Throughout the study period the same technique was applied to estimate the number of drug users in the city and reporting requirements for acute hepatitis B remained the same. We therefore consider it unlikely that the marked decrease in acute hepatitis B observed in drug users is an artefact, i.e. the result of a change in these procedures. From the United Kingdom a similar decrease in acute hepatitis B incidence in drug users since 1985 has been reported.[17]

In conclusion, we found a modest decrease in HIV incidence among IVDUs participating in our follow-up study which correlates with the earlier reported behavioral change in this group.[3] The decline in acute hepatitis B incidence is a convincing indicator of effective behavior change. The largest drop in hepatitis incidence appeared before 1987, a period about which we have very limited data on HIV prevalence and incidence. Possibly a parallel reduction occurred in HIV incidence in that period that we were not able to measure. Nevertheless, recent HIV incidence rates still appear to be disturbingly high, indicating that transmission among this group is continuing. ☐

Acknowledgments

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ANA Releases Draft Report for Health Care Reform

After nearly two years of effort devoted to the development of a reform plan for health care delivery in the nation, the American Nurses' Association (ANA) has released a working draft of the report "Health Care Reform: An Agenda for Action" for review and comment. ANA President Lucille Joel describes the plan as achievable, realistic and cost-effective.

ANA expects to release the report to the public early in 1991, after study by the association's 53 constituents, elected and appointed officials, participants in the ANA Nursing Organization Liaison Forum, and the Tri-Council for Nursing. In this way, ANA expects that the report will reflect an agenda on behalf of the entire nursing profession.

Joel says, "Nursing's plan for health care reform represents a massive effort on behalf of the entire profession to clearly lay out what the nation's nurses believe to be a plan for the delivery of health care services that best meets the needs of the public." The plan offers solutions to many problems affecting the US health care system, including access to care; the lack of or misuse of appropriate providers, settings and delivery arrangements; equitable and affordable financing; and stronger cost controls.

Recognizing that cost and revenue estimates are critical components of health care reform and that adequate funding sources must be available, the ANA plan provides an example for financing. Unique characteristics proposed in the working draft include:

- federal standards to assure access by all to basic levels of health services through public and private health service plans;
- implementation of the plan through incremental steps to be completed by the year 2000 with an initial focus on pregnant women as well as infants and children;
- renewed emphasis on schools, places of work, the home, and other community settings for the delivery of care; and
- reliance on multiple payors—individuals, employers and government—to finance the reformed system.

ANA's Board of Directors planned to review comments received from the nursing community at its December meeting. After approval by the board the plan will be prepared for wide dissemination. For further information, contact ANA, 2420 Pershing Road, Kansas City, MO 64108.

Predictors of Successful Aging: A Twelve-Year Study of Manitoba Elderly

Noralou P. Roos, PhD, and Betty Havens, MA

ABSTRACT

In Manitoba, Canada, a representative cohort of elderly individuals ages 65 to 84 (n = 3,573) were interviewed in 1971 and the survivors of this cohort were reinterviewed in 1983. This analysis assesses the determinants of successful aging—whether or not an individual will live to an advanced age, continue to function well at home, and remain mentally alert. Over 100 separate indicators of demographic and socioeconomic status, social supports, health and mental status in 1971 were available as potential predictors of successful aging. Indicators of access to health care over the period 1970–82 and indicators of diseases over this period were also available as predictors.

Those who aged successfully were shown to have greater satisfaction with life in 1983 and to have made fewer demands on the health care system than those who aged less well. Despite the large number of potential predictors of successful aging which were examined, only age, four measures of health status, two measures of mental status, and not having one's spouse die or enter a nursing home were shown to be predictive of successful aging. (*Am J Public Health* 1991; 81:63–68)

Introduction

The elderly are not a homogeneous group. Some are beset by multiple chronic and/or acute conditions and spend considerable time in hospital. Still others, especially women, live a very long life but are afflicted by increasing frailty or declining mental status, passing their last years in nursing homes. Some elderly manage to avoid all of these scenarios, maintain themselves independently in the community and are described as "aging successfully."[1,2]

While other investigators have attempted to study successful aging, our analysis is unusual in its longitudinal design, its large representative sample, and in focusing on successful aging of an already elderly cohort. It addresses three questions:

1. What proportion of an elderly cohort will age successfully?
2. What are the health care expenditure patterns associated with successful aging?
3. Which characteristics of individuals predict successful aging? (From among all those interviewed in 1971, who will successfully age? And among those who survive to 1983, what are the factors which predict successful aging?)

Methods

In 1971 a large (N = 3,573) representative sample of individuals ages 65–84 living in the community was interviewed as part of the Manitoba Longitudinal Study on Aging.[3,4] In 1983 and 1984 survivors of this cohort were reinterviewed. In addition, the health care expenditures of this cohort over the period 1970 through 1983 were available from administrative records of the universally insured provincial health care program.

The sample was selected as follows: 2.5 percent of the residents of urban areas and 5 percent of the residents of nonurban areas. Analyses were reweighted by the sampling proportions within strata to reflect population figures. A place of residence variable (urban/not) was included in each multivariate model to control for sampling fraction.

Individuals were excluded from the analysis if they were not resident in the province (as judged by coverage in the health care system) for the period of the study, 1970 through 1983 or until death (310); if they were alive and resident in the province in 1983 but not reinterviewed (147); or if the 1971 or 1983 interview was conducted entirely through a proxy (101). (Proxy interviews were used primarily where individuals were too ill or confused to respond.) Native Manitobans (90) were also excluded from the analysis. The final sample for this analysis included 2,943 elderly. A comparison of baseline characteristics of those included and excluded in the study found no difference in age and sex or self-reported health status. However, the interviewer was significantly

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Address reprint requests to Noralou P. Roos, PhD, Professor, Department of Community Health Sciences, University of Manitoba, S101–750 Bannatyne Avenue, Winnipeg, Manitoba R3E OW3, Canada. Dr. Roos holds a Career Scientist Award (6607-1001-48) from NHRDP and is an Associate of the Canadian Institute for Advanced Research; Ms. Havens is Assistant Professor at the University of Manitoba, Provincial Gerontologist, and Acting Assistant Deputy Minister of Community Health Services, Manitoba Department of Health. This paper, submitted to the Journal November 27, 1989, was revised and accepted for publication May 29, 1990.

more likely to rate the respondent's mind as steady or strong in general for those retained in the study.

Successful aging was defined using the 1983 reinterview as follows:

- Alive in 1983.
- Not resident of a nursing home in 1983.
- Did not receive more than 59 days of home care services in 1983.
- In the 1983 interview:

Rated health excellent to fair.
Not dependent in any activities of daily living (getting in and out of bed, dressing, bathing, eating, using stairs).
Did not use a wheelchair.
Did not need help in going out doors.
Able to walk outdoors.
Scored seven or more correct answers on Mental Status Test.

Interviewer judged respondent's state of mind to be steady and strong or a bit weak only after some interview time had elapsed.

Cause of death was taken from death certificates. Less than 1 percent of the individuals (N = 26) died of accidents, poisonings, or violence and therefore lost the chance to age successfully.

Claims data have been shown to provide a highly accurate, reliable, and valid representation of health care use.[5,6] Claims were used to estimate health care expenditure patterns associated with healthy aging over the 14-year period 1970–83 (or until death). The following types of utilization were tracked: days spent in hospital (with days spent in intensive care treated separately); days spent in nursing home; days enrolled in the home care program; all physician visits and surgical fees (for both inpatient and outpatient procedures).

Since nursing home services in Manitoba have been universally insured only since 1973, costs incurred by nursing home residents prior to July 1973 were estimated as described elsewhere.[7] Since September 1974, home care services in the province have also been universally provided at no charge to consumer.

Manitoba hospitals and nursing homes are funded on the basis of global budgets.[8] An estimate of mean costs per day in hospital ($322) and nursing home ($49) was calculated by dividing total Manitoba Health Services Commission payments (including capital repayment) to hospitals and to nursing homes in fiscal 1984 by the number of days of care produced that year by these institutions. The cost of a day in intensive care was estimated to be $1,610, five times the cost of a hospital day. Although per diem costs escalated markedly in both the hospital and nursing home sectors over the time period studied, constant dollar costs were used. The estimate of home care costs per day ($4) were derived by dividing service costs per user over a six-month period by the 182 days in that period.[9] Total payments for physician visits and consultations were divided by the total numbers of visits to obtain an average cost per visit ($17). All costs are in constant 1984 Canadian dollars.

Variables available from the 1971 interview represent potential predictors of successful aging, while diagnoses made during physician visits over the period 1970 through 1982 and variables constructed from administrative records recording events occurring to spouse (death or admission to nursing home) during this same period are potential factors associated with successful aging. Thus predictors of successful aging may include characteristics of the individual's demographic, ethnic and cultural background, socioeconomic characteristics, and characteristics of the support network. The individual's mental status was also assessed in 1971, as well as satisfaction with life, and several characteristics of health status. Factors studied for association with successful aging include changes in the individual's support network over time (whether the individual's spouse predeceased him or her and whether the spouse entered a nursing home prior to the interviewee's death or 1983).

The extent to which specific diseases pose threats to successful aging was examined using both self-reports in 1971 and claims-derived measures of physician contact for five conditions judged to place an elderly individual at "risk for recovery."[10] The Manitoba Cancer Registry was used to determine whether the individual had been diagnosed as having cancer (other than skin). Whether individuals had regular contacts with physicians (and therefore would be available for the early detection and treatment of disease) was measured by the percent of years the individual was alive between 1970 and 1982 during which one or more (or two or more) visits were made to physicians. (See Appendix for a complete listing of predictor variables.)

Two dependent variables were examined: successful aging (yes-no); and successful aging versus survival to 1983 in a dependent state.

In the analyses, using the first dependent variable, all individuals sampled were used to determine which variables were predictive of successful aging. The second analysis examined the association between all dependent variables and successful aging among those who survived to 1983.

Because of the large number of independent variables, a two-step process was followed for each group of variables (see Appendix for grouping). First, the relation of each variable to each dependent variable was tested by univariate analyses. A forward step-wise fitting algorithm for the multiple logistic regression model was used to determine which variables among each group of predictor variables made a contribution to each model at a p value level of 0.05. Age and sex were included in every model. From among those variables within a set which were substantially correlated with one another (i.e. absolute value of Pearson correlation coefficient .5 or greater) only one was selected for modeling. The variables which made the most contribution to model fit, or, where there was little difference, the variables most commonly used in previous studies were selected.

A final multiple logistic regression model was fit including the variables from each set selected as described above. Contributions to the model of the combined effect of 1) sex and income, 2) sex and spousal death/nursing home entry, 3) income and spousal death/nursing home, were assessed by including interaction terms in the model.

The data were randomly split with one-half of the records used to develop the models and to estimate the coefficients of the logistic regression models.[11] The final model was tested on the other half of the records to see if the same variables entered the model and whether the coefficients were estimated in similar magnitude. All data have been pooled for presentation since the larger sample provides more stable estimates of the coefficients.

Results

By our definition, 20 percent of those individuals ages 65 to 84 interviewed in 1971 were judged to have successfully aged in 1983, 22.6 percent were alive but dependent, and 57.5 percent were deceased (Table 1). Males were more likely to be deceased by 1983 than females. Although females ages 65 to 74 were somewhat more likely to have aged success-

TABLE 1—Characteristics of Individuals in 1983 by Age in 1971

	Age in 1971				
	65–74		75–84		
Characteristics of Individual in 1983	Male %	Female %	Male %	Female %	All %
Has successfully aged	24.0	31.5	7.0	5.9	20.0
Alive but dependent	19.9	29.8	12.3	24.4	22.6
Deceased	56.1	38.7	80.7	69.8	57.5
Number (N)	908	930	521	584	2943

TABLE 2—Satisfaction with Life in 1983 in Those Defined as Successfully Aged versus Those Alive but Dependent

	Characteristics of Individuals in 1983		
Satisfaction	Aged Successfully	Alive but Dependent	N
In 1983, how would you describe your satisfaction with life in general at present?			
Excellent	69.6	30.4	182
Good	60.5	39.5	556
Fair	46.9	53.1	193
Poor-Bad	18.1	81.9*	40

*Mantel Haenszel Trend Test: p < .001

TABLE 3—Total Health Care Expenditures over Period 1970–83 by Individuals' Differing 1983 Health Status

	Characteristics of Individuals in 1983			
	Aged Successfully	Alive but Dependent	Deceased	All
Total Health Resource Use* 1970–83 or to Time of Death (%)				
0–$10,000	67.7	30.5	23.5	33.9
10–25,000	20.7	18.4	22.1	21.0
25–50,000	9.5	19.6	21.8	18.8
50–100,000	2.0	14.0	17.0	13.3
$100,000 or more	0.1	17.5	15.7	13.0
Mean Resource Use per Individual (in $ over Period 1970–83 or to Time of Death)				
Total use ($)	10304	53906	50643	43354
Use per year lived ($)	736	3850	5955	4441
Number	583	666	1694	2943

*Resource use is estimated in 1984–85 Canadian dollars as follows: $1,610 per day in intensive care, $322 per hospital day, $49 per nursing home day, $17 per physician visit, $4 per day on home care, and actual $ amount for surgical procedures.

fully than males, there was little difference in successful aging by sex in the older age group (p = .466). Recalculating the percentages in Table 1 (not shown) and focusing upon those who survived to 1983 shows no difference in the percentage of females versus males in the youngest group who successfully aged (p = 0.172). However, among survivors in the oldest age group, males were more likely to successfully age than were females (32.7 percent versus 18.0 percent, p = 0.006).

Remaining independent (the definition of successful aging) is associated with a higher level of satisfaction with life in older age (see Table 2). Separate comparisons by age confirmed this relationship (data available on request to author). However, maintaining functional independence is not a guarantee for being satisfied with one's life in advanced old age. Moreover, losing one's independence is not necessarily judged a disaster. Among those reporting their satisfaction with life to be excellent, 30.4 percent were classifed as "alive but dependent" in 1983.

Individuals who remained independent made markedly fewer demands on the health care system (Table 3). About two-thirds (67.7 percent) of those aging successfully used less than $10,000 worth of health care resources over the 14-year period 1970 through 1983, averaging $736 per year. In contrast, those who lived to 1983 in a more dependent state averaged $3,850 per year over the 14-year period. Those who died before 1983 had the highest resource use averaging $5,955 per year lived. Further analyses (not presented in Table 3) demonstrated that the low resource utilization patterns of those aging successfully were observed both for males ($768 per year lived) and females ($710) and were independent of age ($744 per year for those aged 65–74 in 1971 and $671 for those aged 75–84).

Table 4, using multivariate analysis, suggests that the most important factors distinguishing those who successfully aged from all others interviewed in 1971, apart from age, were: bad outcomes occurring to a spouse, self-reported health in 1971, retirement because of poor health, not developing cancer or diabetes, and not having good mental status. Females were no more likely to age successfully than were males.

Factors which distinguish those who successfully aged from those who survived to 1983 but were functionally dependent are shown in Table 5. After controlling for age, self-rated health was again a strong predictor of successful aging. Individuals who did not have diabetes were also more likely to age successfully, although cancer had no similar impact. The odds of successful aging were decreased if one's spouse died or was assessed for or placed in a nursing home.

Discussion

Successful aging has been defined in terms of an individual retaining the ability to function independently.[1,2] Those who age successfully remain out of institutions and do not have continuing input from a

TABLE 4—Adjusted Relative Odds of Successful Aging as Estimated by the Multiple Logistic Regression Model. I. Comparing Those Who Age Successfully with Those Who Died or Became Functionally Dependent by 1983

	Relative Odds Estimate	95% Confidence Interval
Age of Individual in 1971		
65–69/80–84	28.8	(13.9, 59.7)
70–74/80–84	12.7	(6.1, 26.3)
75–79/80–84	4.5	(2.1, 9.7)
Sex		
female/male	1.2	(1.0, 1.5)
Spouse Assessed/Placed in Nursing Home (1971–82)		
yes/others	0.2	(0.1, 0.5)
Spouse Died 1971–82		
yes/others	0.4	(0.2, 0.6)
Self-Rated Health 1971		
excellent/poor bad	7.0	(4.0, 12.4)
good/poor bad	4.0	(2.4, 6.8)
fair/poor bad	2.4	(1.4, 4.2)
Reason for Retiring (1971)		
poor health, compulsory/others	0.6	(0.5, 0.8)
Cancer (1970–82)		
no/yes	3.5	(2.4, 5.1)
Had Physician Visit(s) for Diabetes (1970–82)		
no/yes	1.7	(1.3, 2.5)
Interviewer Assessed State of Mind (1971)		
steady and strong/else	2.5	(1.4, 4.6)
Score on Mental Status Index (1971)		
High score (9 or 10)/Lower score (0–8)	1.6	(1.1, 2.4)

Model Chi-Square = 514 with 15df. Region of residence was included in the model to control for sampling fraction. Logistic regression results were based on comparing 583 individuals who aged successfully with 2,350 individuals who died or became functionally dependent by 1983. The score on the mental status index, not having diabetes, and rating one's health as fair, reached the 0.05 level of significance in only one of the split half models and the final model.

home health agency. They remain mobile and competent in all the activities of daily living. While these individuals defined their health status in 1983 as fair or better, we deliberately did not label these individuals the "healthy aged." Healthy implies the absence of disease and further, individuals with the "same disease" may have a different sense of well-being and ability to function.[12]

Our results are interesting not only because of the variables they include, but also for the variables which they do not include. None of the measures of socioeconomic status were significant in the final models, even though the measures available in 1971 were extensive. Also, an index combining high scores on the education, occupation and income variables was constructed but did not prove significant.

Two recent studies[13,14] have reported that low incomes are associated with poor health outcomes for Canadians despite universally insured medical and hospital care. Guralnik and Kaplan[12] also found income predictive of a high level of functioning in the 19-year follow-up of Alameda County residents. Two factors may help to account for our contradictory findings compared to these earlier studies. The income effect may decrease with age because: not as many poor people as others survive long enough to be classified as "old"; the income gradient among the elderly may narrow with government transfer payments improving the lot of the poor, while advancing age is depleting the resources of the wealthy. Haan, *et al*,[15] compared age-specific mortality rates among White Alameda County residents of poverty and non-poverty areas. While in the younger age groups the mortality rates in poverty areas were much higher, among those ages 65 and older, there was essentially no difference. Finally, the one group in Canadian society which is most economically disadvantaged, Native Canadians, were excluded from the analysis.

This study presents strong evidence that regular contact with physicians, even among elderly individuals (one of the highest risk groups in the population), is not related to successful aging. Regular contact with the health care system was measured with a high degree of accuracy. Neither the percentage of years in which individuals made one or more visits, nor the percentage of years in which two or more visits were made, was associated with successful aging.

Many measures of disease and functional status were not useful predictors of successful aging. None of the measures of poor functional status in 1971, including indicators that an individual needed help in the activities of daily living or in instrumental activities, none of the self-reported health conditions (including stroke, arthritis, etc.) and only one of the conditions for which physician visit data were available (diabetes) was negatively associated with successful aging.

These findings are contrary to most preconceptions and to some literature. The National Institute of Mental Health's Human Aging II Study,[16] a longitudinal follow-up of a small number of very healthy males, found even asymptomatic disease was negatively predictive of longevity. However, that study focused on longevity alone and longevity is only one (albeit an important) aspect of successful aging.

Having cancer was negatively associated with successful aging but primarily influenced longevity; people with cancer were less likely to live long enough to age successfully by our criteria. However, if such individuals did survive to 1983 (6.7 percent of those surviving to 1983 had cancer), they were just as likely to have aged successfully as those without cancer.

Since age is strongly related to mortality, odds ratios for categories associated with successful aging are greater for the model whose comparison group included decedents (Table 4). The magnitude of the odds ratio for the female sex decreased somewhat from model one to model two (Table 5) reflecting the fact that, although more females may survive, their chance of aging successfully appears to be lower than that for males who survive.[17] A comparison of self-rated health in the two models suggests the distribution of self-rated health in 1971 among those who subsequently died was similar to that among those who remained alive but dependent.

Previous work in Manitoba demonstrated that self-rated health was a predictor of mortality over a six-year period independent of objective health status[18] and this relationship has been confirmed by

TABLE 5—Adjusted Relative Odds of Successful Aging as Estimated by the Multiple Logistic Regression Model. II. Comparing Those Who Age Successfully with Those Who Were Alive in 1983 but Dependent

Specification of Variable	Relative Odds Estimate	95% Confidence Interval
Age of Individual in 1971		
65–69/80–84	11.5	(5.3, 25.0)
70–74/80–84	5.6	(2.6, 12.3)
75–79/80–84	2.7	(1.2, 6.2)
Sex		
female/male	0.8	(0.6, 1.0)
Spouse Assessed/Placed in Nursing Home (1971–82)		
yes/others	0.1	(0, 0.4)
Spouse Died 1971–82		
yes/others	0.4	(0.3, 0.8)
Self-Rated Health 1971		
excellent/poor bad	5.5	(2.9, 10.6)
good/poor bad	3.0	(1.7, 5.5)
fair/poor bad	2.2	(1.2, 4.0)
Reason for Retiring (1971)		
poor health, compulsory/others	0.7	(0.5, 0.9)
Cancer (1970–82)		
no/yes	1.0	(0.6, 1.7)
Had Physician Visit(s) for Diabetes (1970–82)		
no/yes	1.5	(1.0, 2.2)
Interviewer Assessed State of Mind (1971)		
steady and strong/else	1.9	(1.0, 3.9)
Score on Mental Status Index (1971)		
High score (9 or 10)/Lower score (0–8)	1.9	(1.2, 2.8)

Model Chi-Square = 208 with 15df. Included in the model was region of residence to control for sampling fraction. Logistic regression results based on comparing 583 individuals who aged successfully with 662 individuals who were alive in 1983 but dependent. Age 75–79, having a spouse who died, retiring because of poor health, having a visit for diabetes, reporting fair health and having a high score on the mental status index, reached the 0.05 level of significance in only one of the split half models and the final model.

others.[19] The present study extends these findings by demonstrating that self-rated health is not only an important predictor of mortality over the longer 12-year period, but is also one of the few factors associated with successful aging among those who survive. Muller[20] suggested that the predictive contribution of self-rated health could be partly due to Mossey and Shapiro's including indicators developed from utilization data in their measure of objective health status. She suggested pessimism and depression might be the actual causal factors. Although measures of life satisfaction in 1971 were predictive of successful aging in early models, they made no contribution here to the final models which included the self-rated health measure.

There were measures unavailable to this project which might have made useful contributions. The most obvious limitation of the data was the lack of biological measures of function (blood pressure, etc.) and the lack of measures of health practices (smoking history, alcohol consumption, regular physical activity, etc. although the latter could be approximated by the questions pertaining to leisure time activities). The Alameda County study found individuals with more low risk practices to have lower mortality rates.[21] The Framingham Study, after 21 years of biennial observations, also reported that alcohol intake, smoking, ventricular rate, and education were related to good functioning status for men although, for women, the only significant predictor other than age was education.[22]

In conclusion, a significant group of elderly do age successfully. These individuals express more satisfaction with their lives and incur substantially fewer health expenditures than other elderly. Individuals at particular risk of not aging successfully include those with poor self-assessed health, whose spouse has died, whose mental status is somewhat compromised, who developed cancer, and those who are forced to retire or retire because of poor health. ☐

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APPENDIX

Variables Available to the Analysis
(All Items from 1971 Interview Except as Indicated)

- **Ethnicity and Cultural Characteristics of Individual**
  - National descent
  - Language spoken
  - Date of arrival in Canada
  - Language of interview

- **Socioeconomic Characteristics of Individual**
  - Level of education attainment (number of years or grades completed in school)
  - Major occupation when employed
  - Current employment in major or other occupation
  - Major reason for continuing working (if applicable)
  - Total monthly income (a total of income reported from three sources: own resources—wage, rents, interest, etc.; pensions or allowances; other sources—children, etc.)
  - Disposable income (calculated by subtracting monthly expenses from monthly income)
  - Percent of income from private sources
  - Whether income satisfied current needs
  - Whether income would satisfy future needs
  - Would extra income be spent on food, medical needs, etc.
  - Ownership of current residence
  - Characteristics of residence
    - Heating satisfactory, lighting satisfactory, complete bathroom available, bathroom sharing, a room of one's own?
    - Residence in a senior citizens' housing unit

- **Characteristics of Individual's Social Network**
  - Marital status
  - Living arrangements (with spouse, with spouse and others, other)
  - Number of relatives, amount of contact with them and "nearness"
  - Frequency of contact with neighbors, friends
  - Death of spouse between 1971 interview or person's death or December 31, 1982*
  - Spouse assessed for or entered nursing home during same period*

- **Individual's Mental Status**
  - Score on 10 question mental status test[23]
  - Frequency of forgetting names of friends or relatives, time of day. How quickly recalled?
  - Interviewer's assessment of respondent's:
    - —state of mind (steady and strong to weak or unsteady)
    - —attitude (friendly to hostile)
    - —comprehension of questions (satisfactory to unsatisfactory)

- **Satisfaction with Life**
  - In general
  - Score on 20 question Life Satisfaction Index "A"[24]

- **Individual's Health Status**
  - Self-reported health: "For your age, would you say your health is excellent, good, fair, poor or bad?"
  - Reporting any of 15 health problems including diabetes, chest problems such as asthma or emphysema
  - Under treatment for any of these conditions
  - Amount of time in bed at home because of illness in last year: none, less than one week, one week to a month, more than a month
  - Help required in basic activities of daily living (getting in and out of bed, dressing, bathing, eating, using stairs)
  - Number of activities requiring help from someone outside the house (light housework, making tea, preparing a hot meal, shopping, laundry, heavy housework, financial affairs)
  - Number of activities requiring help from someone inside or outside the house (nursing care, cutting toenails, going out of doors, watching TV or reading)
  - Required attention of physician or nurse when last sick
  - Moved because of ill health
  - Retired because of poor health
  - Living in a senior citizen housing unit (individuals tend to move to these rental units when they have difficulty maintaining themselves at home usually for health and/or economic reasons)
  - Unmet health care needs for eyeglasses, hearing aid, false teeth, cane, special shoes, wheelchair or medicine
  - Mean number of physician visits (per year alive) for conditions identified as increasing an individual's "risk of not recovering from an illness"[10] e.g. angina, arthritis, arrhythmia, diabetes, hypertension over period 1970–82*
  - Number of above listed conditions for which made one or more physician visits (per year alive); for which made two or more physician visits (per year alive)*
  - Cancer (excluding skin cancer) diagnosed 1970–82**

- **Participation in Leisure Time Activities**
  - 21 alternatives including visiting, telephoning, listening to radio or TV, work, etc.
  - Which activities are most important

- **Access to Care**
  - Difficulty getting to see a doctor/dentist because of transportation, language problems or distance
  - Percentage of years alive in which one or more visits to physician were made; in which made two or more visits, 1970–82*

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*Indicator derived from claims files of Manitoba Health Services Commission.
**Indicator derived from files of the Manitoba Cancer Registry.

Does Maternal Tobacco Smoking Modify the Effect of Alcohol on Fetal Growth?

Jørn Olsen, MD, PhD, Altamiro da Costa Pereira, MD, and Sjurdur F. Olsen, MD

ABSTRACT

Smoking and drinking habits were registered by a self-administered questionnaire in 36th week of gestation in 11,698 pregnant women, more than 80 percent of all such women in two Danish cities 1984–87. Alcohol consumption of 120 g/week or more was associated with a greater reduction in the average birthweight in the babies of smokers than of non-smokers (about 40 grams for the non-smokers and about 200 grams for the smokers). This is particularly striking considering that the average birthweight for smokers is lower than for non-smokers. A birthweight difference of more than 500 grams was found between babies of mothers who neither smoked nor drank and mothers who smoked and drank heavily. Our data suggest that women's smoking habits should be taken into consideration when giving pregnant women advice about drinking. (*Am J Public Health* 1991; 81:69–73)

Introduction

It has been known for many years that smoking during pregnancy reduces fetal growth[1] and it is also well established that growth retardation is part of the fetal alcohol syndrome.[2] Tobacco smoking relates to fetal growth and birthweight in a dose-response manner[1] but the association with alcohol consumption remains less clear. In a recent publication,[3] only alcohol intake of 120 grams or more per week was associated with low birthweight, and it was suggested that health education should be directed mainly toward mothers who drink 100 grams of alcohol or more per week. Another recent study[4] found that alcohol had an effect on birthweight only in smokers.

The purpose of our study was to investigate whether it is reasonable to assume a threshold for the effect of alcohol on birthweight and other growth parameters according to smoking habits.

Methods

Detailed information was collected on drinking and smoking habits of pregnant women attending the midwife centers in Odense and Aalborg, Denmark during their 36th week of gestation, from April 1984 to April 1987. The mothers sent a completed questionnaire directly to the Department of Social Medicine, Odense University. The medical records and questionnaires were given a study number which was not related to the personal identification number. After delivery, obstetrical information was extracted from the medical records and linked to the questionnaire data by means of the study number.

More than 95 percent of all pregnant women in two defined geographical areas at the 36th week of gestation were covered, for a total of 13,815 women. After excluding multiple births, data were available from the questionnaires and medical files for 11,850 women. Most of the non-responders failed to answer reminders, and some left the study area before delivery so that no data were available regarding the child. Also excluded were 152 women who gave birth before the 36th week of gestation; the data file thus consisted of 11,698 women.

Detailed questions were asked concerning beer, wine, spirit, coffee, and tea consumption before and during pregnancy. The number of alcoholic drinks per week was calculated; each drink was considered to contain 12 grams of absolute alcohol according to normal quantities of alcoholic beverages per glass (or bottle of beer) in Denmark. The data on smoking and drinking were self-reported and referred to average consumption during the first 36 weeks of gestation.

Data on gestational age, weight increase during pregnancy, placental weight and measurements of the child (weight, length and head circumference) were recorded from the mothers' and infants' hospital records. All were measured blind to knowledge of questionnaire data. All deliveries took place at one central hospital in each region. The midwives who car-

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Address reprint requests to Jørn Olsen, MD, PhD, Institute of Social Medicine, University of Aarhus, Høegh Guldbergs Gade 8, 800 Aarhus C, Denmark. Dr. S. Olsen is also at that address: Dr. Pereira is with the Department of Hygiene and Epidemiology, Faculty of Medicine, Porto, Portugal. This paper, submitted to the Journal February 26, 1990, was revised and accepted for publication August 28, 1990.

ried out the recording had a comparable training and conducted the assessments in a standardized way. The study was part of a large community trial among pregnant women in Denmark.[5]

In the data analysis, multiplicative as well as additive models have been used (logistic regression and analysis of variance). The variables which were most closely associated with alcohol consumption and which were known to be predictors of fetal growth were included as confounders.

Results

Table 1 displays the characteristics of the participating women. Smoking and drinking habits during the first 36 weeks of gestation are shown in Table 2. Thirty-five percent were non-smokers who drank 1–29 grams of alcohol per week. Almost 200 women reported that they consumed 120 grams (10 drinks) or more per week and 832 indicated a smoking history of 15 cigarettes or more per day. The proportion of newborns with a birthweight of less than 2500 grams varied widely with a tendency to correlate with the frequency of alcohol and tobacco use.

Most newborn measures decrease with increasing cigarette and alcohol intake. Among smokers the decrease associated with alcohol is greater than among non-smokers (see Table 3).

In Table 4, the results of the analysis of variance of the data shown in Table 3 are presented. Among non-smokers, alcohol consumption was found to be statistically significantly associated with placental weight and birthweight (test for trend). Among smokers of 1–9 cigarettes, alcohol was associated with birth length, and in the subgroup of heavy smokers alcohol was significantly associated with all measures (except birth length).

All of the variables in Table 4 depended upon the duration of pregnancy and in two additional analyses of variance, confounding by gestational age was addressed (data not shown). In one model, gestational age was included as a covariate and, in another model, the analyses were restricted to women who gave birth between the 39th and 41st weeks of gestation. In these analyses, the adjusted differences from the mean were in the same direction but usually a little smaller. A statistically significant association between alcohol consumption and placental weight was still found for non-smokers. In the group of moderate smokers, significant associations were found for placental weight, birthweight and birth length (but not in the restricted analyses). In the group of heavy smokers, all associations remained statistically significant at the 0.05 level (except in the restricted model with birth length).

Clear "dose-response" relationships were not seen in any models in Table 4, except for birthweight among heavy smokers. However, random variation due to small numbers in the sub strata should be taken into consideration. Among heavy smokers, women who drank more than 30 grams of alcohol per week had a substantially lower birthweight. This is particularly striking considering the generally low mean for that group of women.

Table 5 presents odds ratios for low birthweight according to alcohol consumption compared with frequency of low birthweight for non-drinkers. Each smoking stratum represents a separate analysis. A two- to three-fold increase was seen for all consumers of 120 grams of alcohol or more per week. The marginal values in the table show an excess risk of low birthweight in the group of "heavy drinkers" only. The prevalence of low birthweight increased steadily with increasing cigarette consumption, as can be seen in the right hand column.

All analyses in Table 5 were repeated with the subset of women who gave birth in the 39th–41st weeks of gestation and weight increase during pregnancy was also included in the models. By restricting analyses to deliveries at term, a four-fold increase of low birthweight was found for consumers of more than 120 grams of alcohol per week (data not shown).

After controlling for average alcohol consumption, no association between binge drinking and any of the pregnancy outcome measures was seen (data not shown).

According to the present data set, neither the additive nor the multiplicative model of interaction between alcohol consumption and tobacco smoking on any of these measured outcomes fitted the data very well, despite the fact that both models were acceptable according to formal statistical testing of interaction terms.

Discussion

Birthweight was closely correlated with the combined exposure to alcohol as

TABLE 1—Age, Parity, Gestational Age at Birth and Weight Increase during Pregnancy for 11,698 Participating Women

Characteristics	N. Cases	%
Age (years)		
under 24	3,473	29.7
25 to 34	7,356	62.9
35 or more	869	7.4
Parity		
0	5,628	48.1
1	4,412	37.7
2 or more	1,658	14.2
Gestational age in weeks		
under 39	1,890	16.2
39 to 40	6,122	52.3
41 or more	3,686	31.5
Weight increase in kg*		
under 11	2,923	34.4
12 to 15	2,970	35.0
16 or more	2,599	30.6

*3,206 missing from start of pregnancy to the time of delivery

TABLE 2—Number of Pregnancies According to Maternal Drinking and Smoking Habits: Bottom line, Newborns < 2500 g in percent

Number of Cigarettes per Day	Grams of Alcohol per Week						
	0	1–	30–	60–	90–	120+	All
0	1,267	4,088	1,082	404	103	101	7,045
	2.0	1.6	1.8	1.5	1.0	4.0	
1–4	102	368	110	55	10	15	660
	2.9	3.5	2.7	3.6	0.0	13.3	
5–9	294	839	213	74	26	28	1,474
	5.1	4.2	2.8	1.4	0.0	10.7	
10–14	315	894	218	70	31	24	1,552
	4.1	4.8	7.2	5.7	9.7	0.0	
15+	172	433	132	40	26	29	832
	3.5	6.2	5.3	7.5	11.5	25.0	
All	2,150	6,622	1,755	643	196	197	11,563*

*135 missing cases

TABLE 3—Means of Placental Weight, Birthweight, Birth Length and Head Circumference According to the Mother's Alcohol Consumption during Pregnancy, Stratified by Smoking Status

Smoking Status (Daily)	Alcohol in g/Week	Placental Weight in g	Birthweight in g	Birth Length in cm	Head Circumference in cm
Non-smokers	0	569	3,572	52.6	35.4
	1–	561	3,579	52.6	35.4
	30–	557	3,602	52.8	35.5
	60–	537	3,582	52.6	35.5
	90–	532	3,554	52.8	35.4
	120+	534	3,536	52.4	35.5
Subtotal	Mean	560	3,580	52.6	35.4
	(Stddev)	(132)	(483)	(2.3)	(1.5)
1–4 cigarettes	0	567	3,487	52.2	35.1
	1–	557	3,493	52.3	35.3
	30–	572	3,479	52.0	35.4
	60–	564	3,573	52.6	35.3
	90–	626	3,727	53.4	35.6
	120+	450	3,141	51.3	34.7
Subtotal	Mean	560	3,492	52.2	35.3
	(Stddev)	(135)	(510)	(2.5)	(1.6)
5–9 cigarettes	0	555	3,330	51.6	35.0
	1–	548	3,321	51.5	35.0
	30–	528	3,261	51.5	34.9
	60–	522	3,379	51.9	34.8
	90–	524	3,444	52.1	35.5
	120+	489	3,144	50.6	34.6
Subtotal	Mean	544	3,316	51.5	35.0
	(Stddev)	(121)	(458)	(2.2)	(1.6)
10–14 cigarettes	0	572	3,357	51.5	35.0
	1–	548	3,282	51.4	34.9
	30–	535	3,259	51.4	34.8
	60–	528	3,291	51.4	34.7
	90–	540	3,123	50.5	34.4
	120+	493	3,116	50.9	34.4
Subtotal	Mean	549	3,289	51.4	34.9
	(Stddev)	(132)	(467)	(2.2)	(1.5)
15+ cigarettes	0	546	3,270	51.4	34.9
	1–	561	3,266	51.3	34.8
	30–	531	3,207	51.3	34.6
	60–	524	3,123	50.8	34.6
	90–	539	3,169	50.8	34.8
	120+	472	3,023	50.8	34.0
Subtotal	Mean	548	3,239	51.2	34.7
	(Stddev)	(128)	(459)	(2.3)	(1.5)
Total	Mean	556	3,478	52.2	35.2
	(Stddev)	(132)	(497)	(2.4)	(1.6)

well as smoking, but a number of limitations of the study should be kept in mind. The sample was truncated at 36 weeks, and therefore precluded study of babies born at earlier gestation. Since data on smoking and drinking habits are self-reported, intake of alcohol especially was likely to be underreported, particularly high consumption levels. For this reason, no estimates were made for alcohol consumption of more than 120 grams per week. Few women indicated no drinking at all during pregnancy, which agrees with our impression that Danish women usually report moderate and socially acceptable levels of drinking without hesitation.

It is likely that the heavy drinkers were more prevalent among non-responders than responders, but since exposure data were collected prior to the deliveries, bias is not a serious problem. Nevertheless, the effects shown may be underestimates.

Sulaiman, *et al*,[3] suggested that health education during pregnancy should be directed toward mothers who drink more than 100 grams of alcohol per week since no statistically significant effect was seen below this level. However, that study was rather small with only a few women consuming more than 100 g absolute alcohol per week. Lumley, *et al*,[6] did not detect any harmful effect on fetal outcome of alcohol consumption during pregnancy at levels below two glasses a day. Our data indicate that health education should take smoking habits into consideration when giving advice about alcohol use during pregnancy. If pregnant women do not give up smoking while pregnant, a much lower level of alcohol intake should be aimed at. Furthermore, it must be taken into consideration that only a limited set of potential adverse effects of alcohol or smoking combined have been studied. Martin, *et al*,[7] found that maternal alcohol intake, coupled with smoking, exerted an interactive and deleterious effect upon newborn learning.

Our data to some extent supported the findings that consumption of beer is more closely related to adverse outcomes of pregnancy,[3,9,10] but it is likely that the association was confounded by social factors not recorded in the study (data not shown). We cannot recommend that health education should address the types of beverages differently at present.

We found that the average consumption of alcohol was much more important than binge drinking (data not shown). However, this may not be true for other adverse outcomes of pregnancy. Keeping alcohol's teratogenic effect in mind,[2] it is still wise to warn against binge drinking while pregnant.

Kariniemi, *et al*,[11] found that maternal alcohol consumption only decreased the birthweight among female fetuses. However, no such effect was seen in our study (data not shown).

Low birthweight remains the most important determinant of perinatal mortality and impaired development. When alcohol consumption increases among women, one should be aware of the fact that combined exposure of both alcohol and tobacco may reduce birthweight to a substantial degree, as illustrated by the increased ORs of low birthweight. The difference in birthweight between the lowest level of alcohol and tobacco use and the highest levels was found to be more than 500 grams.

In a large scale study, Mills, *et al*,[12] found somewhat different effects than we did. In our data, heavy alcohol consumption seemed to have a stronger effect on birthweight among smokers than non-smokers, whereas this phenomenon was apparently absent in Mills' data. Like

TABLE 4—Adjusted Differences from Mean Values of Placental Weight, Birthweight, Birth Length and Head Circumferences According to the Mother's Alcohol Consumption during Pregnancy, Stratified by Smoking Status: Analysis of Variance*

Smoking Status (Daily)	Alcohol in g/Week	Placental Weight in g	Birth-weight in g	Birth Length in cm	Head circumference in cm
Non-smokers	0	9	1	−.0	−.0
	1−	2	−2	−.0	−.1
	30−	−1	21	.1	.1
	60−	−29	−14	−.1	.0
	90−	−34	−48	.0	−.1
	120+	−35	−44	−.2	.0
	Mean	565	3,582	52.6	35.4
1–9 cigarettes	0	12	24	.1	.0
	1−	2	1	−.0	.0
	30−	−9	−45	−.1	−.0
	60−	−9	72	.4	−.1
	90−	−3	121	.5	.4
	120+	−65	−245	−1	−.5
	Mean	554	3,377	51.7	35.1
10+ cigarettes	0	13	63	.1	−.0
	1−	5	−9	−.0	.1
	30−	−19	−61	.1	−.2
	60−	−22	−31	−.1	−.1
	90−	−6	−145	.0	−.9
	120+	−73	−206	−.2	−.7
	Mean	554	3,283	51.3	35.4

*All values have been adjusted for parity (0, 1, 2, +), school education (−9y, 10y, 11 + y) and age. Multiple classification analysis, adjusted for main effects and covariates in 3 separate subgroups according to smoking habits.

TABLE 5—ORs (and 95% CI) of Low Birthweight (< 2,500 g) According to the Mother's Smoking and Drinking Habits during Pregnancy (logistic regression)

Smoking* Status (Daily)	Alcohol in g/Week						Adjusted for Alcohol
	0	1−	30−	60−	90−	120+	All
0	1.0	0.8 (0.5–1.3)	0.9 (0.5–1.6)	0.8 (0.3–1.9)	0.5 (0.1–3.8)	2.2 (0.7–6.5)	1.0
1–9 cigarettes	1.0	0.9 (0.5–1.6)	0.6 (0.3–1.4)	0.5 (0.1–1.7)	— —	2.7 (0.9–8.0)	2.2 (1.7–3.0)
10+ cigarettes	1.0	0.8 (0.5–1.3)	0.6 (0.3–1.1)	0.9 (0.4–2.2)	1.6 (0.6–4.1)	2.1 (0.8–5.2)	3.4 (2.6–4.3)
Adjusted for smoking All	1.0	1.0 (0.7–1.3)	1.0 (0.7–1.5)	0.9 (0.5–1.5)	1.1 (0.5–2.5)	2.7 (1.5–4.8)	

*Logistic regression made on those subgroups of women according to smoking habits and on all in the study (marginal row and column) all analyses adjusted for age, school education, and parity. Alcohol and smoking entered the model as "dummy variables."

Mills, *et al*, we found little effect of low to moderate drinking among non-smokers but among heavy smokers reductions in average birthweight tended to appear already at an alcohol consumption of 30–59 g/week, and this pattern was not evident from the Mills data. One problem of comparing the two studies is that they did not use the same standards for alcohol and smoking exposure. More important, however, is the fact that our exposure data were collected at the 36th week of gestation while Mills' data were collected at the first visit in the first trimester. The generally stronger effects of alcohol on birthweight seen in our study may thus reflect the impact of a more relevant time period.

Although the significance of placental weight is disputed, it has been found to be a predictor of growth later in the childhood.[13] It is therefore interesting that alcohol consumption was found to be associated with relatively large reductions in average placental weight, although it should be kept in mind that the effects might have been exerted via fetal growth since birthweight and placental weight are closely associated.

Unfortunately, only some of the potential confounders could be controlled in this study. Nutrition, in particular, may be an important confounder in a study of alcohol consumption and fetal growth. Furthermore, it is likely that heavy drinkers may be less willing to follow prenatal care. Use of other drugs may be a confounder in some countries but it is unlikely in this sample.

The combined effect of alcohol and tobacco smoking should be taken into consideration when advising on health behavior during pregnancy. ☐

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Pew Commission Advances Agenda Action for Nation's Health Professional Schools

The Pew Health Professions Commission met for the second time on November 7 and 8 in Washington, DC to advance 18 reform recommendations as part of a national Agenda for Action that addresses health care and educational issues challenging the nation's health professional schools. This agenda is built on an analysis of the current health care environment and the major forces and trends which will shape health care in the United States over the next 15 years.

The Commission's recommendations for discussion include a redefinition of the core science curriculum; a reorientation of the teacher-learning process; a shift in patient care training setting; a dissolution of the boundaries between the health professions; a redirection of faculty incentives; the creation of new governance and organizational arrangements; and a change in the orientation of the health professions from illness to health.

The Pew Commission will continue to refine the Agenda for Action over the next three months, and is expected to announces its final Agenda for Action in April 1991. The five objectives of the Pew Commission are to: establish a national Agenda for Action; develop appropriate policy recommendations for federal, state, professional and institutional agencies which will support implementation of the Agenda for Action; communicate the Agenda for Action; provide institutional change resources; and explore and fund demonstration projects at health professional schools across the nation.

The Pew Health Professions Commission is an outgrowth of the Future of the Health Professions project, a program of Duke University and The Pew Charitable Trusts. Over the past six years, The Pew Charitable Trusts has invested over $60 million of programming support in its Strengthening the Health Professions initiative and has secured an additional $40 million of leveraged support from other foundations for this initiative. For further information, contact Pew Health Professions Commission, Duke University Medical Center, 3101 Petty Road Suite 1106, Durham, NC 27707. Tel: (919) 490-6695; Fax (919) 489-5907.

Admission and Mid-Stay MedisGroups® Scores as Predictors of Death Within 30 Days of Hospital Admission

ABSTRACT

We examined the ability of MedisGroups, a severity measure based on clinical data abstracted from the medical record, to predict mortality 30 days following admission. MedisGroups measures severity both on admission and approximately one week into the hospital stay. The data base was a random sample of 20,985 admissions of Medicare beneficiaries with one of six conditions from 833 hospitals in seven states between January 1985 and May 1986. In all six conditions, higher admission and mid-stay severity scores were generally associated with higher risk of death. Across the six conditions, the R-squared values for predicting death using admission scores ranged from 0.01 to 0.16; R-squared values using mid-stay scores ranged from 0.03 to 0.34; and R-squared values from combining admission and mid-stay scores ranged from 0.05 to 0.41. Admission MedisGroups score was thus only modestly predictive of 30-day mortality. While the mid-stay score was more powerful, it may not be an appropriate severity adjuster to screen for quality using hospital mortality rates because it could be influenced by substandard care. *(Am J Public Health* 1991; 81:74–78)

Lisa I. Iezzoni, MD, MS, Arlene S. Ash, PhD, Gerald Coffman, MS, and Mark A. Moskowitz, MD

Introduction

The Federal Government annually publishes mortality rates of almost 6,000 United States hospitals caring for Medicare beneficiaries. Although release of the mortality data is accompanied by numerous and appropriate disclaimers about their utility, the data are clearly intended to serve as an initial clue about the quality of hospital care, especially in an era of tightening cost constraints. The first hospital mortality statistics, published in March 1986, indicated that 142 hospitals had significantly higher death rates than predicted and 127 had significantly lower rates.[1] However, the facility with the most aberrant death rate (87.6 percent of Medicare patients dying compared with a predicted 22.5 percent) was a hospice caring for terminally ill patients. The government model had not adequately considered severity of illness.

This incident prompted a flurry of research activity, which has resulted in more sophisticated and methodologically sound approaches for producing the Medicare hospital mortality statistics.[2] Since 1987, the model has adjusted for patient age, secondary diagnoses, and a variety of other characteristics available in Medicare's administrative files; in addition, to control better for case mix and to accommodate the possibility that hospitals may perform better in some clinical areas than in others, condition-specific hospital mortality rates are also produced for selected diseases. However, data limitations remain. The information contained in Medicare's administrative files is confined to basic demographic, diagnostic, and procedural data not considered adequate to the task of severity adjustment by some critics.[3–7] The purpose of this research was to examine the performance of MedisGroups®, a severity measurement system based on data drawn from the medical record, as a predictor of mortality.

Methods

Description of MedisGroups

MedisGroups assesses severity as the risk of imminent organ failure independent of diagnosis, using information on key clinical findings (KCFs) abstracted from the medical record. MedisGroups is a registered trade name; the system is owned and marketed by MediQual Systems, Inc., in Westborough, Massachusetts. Severity is scored on admission on a five-point scale ranging from 0 to 4. For patients who remain hospitalized for the required period, a second severity rating is performed, the so-called "mid-stay" review, generating an appraisal of extent of "morbidity" (no morbidity, morbid, or major morbidity). MedisGroups is described in detail elsewhere,[8–11] and more information on the system is presented in an appendix. This description and the data used for this study pertain to the 1987 version of MedisGroups.

Address reprint requests to Lisa I. Iezzoni, Division of General Medicine and Primary Care, Department of Medicine, Beth Israel Hospital, 330 Brookline Avenue, Boston, MA 02215, Tel: (617)735-4476. Dr. Iezzoni is also responsible for negotiations concerning this manuscript. All the authors were with the Health Care Research Unit, Section of General Internal Medicine, Evans Memorial Department of Clinical Research and Medicine, Boston University School of Medicine. Dr. Iezzoni was also with the Health Policy Institute, Boston University School of Public Health. This paper, submitted to the Journal January 10, 1990, was revised and accepted for publication June 12, 1990.

TABLE 1—Number of Cases by Admission MedisGroups Score and Percent Dying within 30 Days by Condition

Condition	Admission MedisGroups Score 0	1	2	3	4	Total
	Number of cases (percent dying within 30 days)					
Pneumonia or COPD	73 (1.4)	656 (2.9)	1,985 (9.0)	798 (28.7)	40 (67.5)	3,552 (12.8)
Coronary revascularization	1,073 (2.5)	1,228 (4.7)	782 (6.9)	273 (19.1)	11 (36.4)	3,367 (5.8)
Acute myocardial infarction	159 (6.3)	536 (8.2)	1,516 (15.4)	974 (40.9)	127 (80.3)	3,312 (23.8)
Congestive heart failure	119 (5.9)	396 (7.6)	1,792 (10.3)	1,143 (23.0)	74 (55.4)	3,524 (14.9)
Cholecystectomy	1,002 (1.2)	1,094 (1.4)	1,192 (2.8)	298 (5.4)	3 (33.3)	3,589 (2.2)
Prostate surgery	1,258 (0.6)	1,769 (0.8)	557 (2.5)	56 (7.1)	1 (0.0)	3,641 (1.1)

When MedisGroups is used for quality assessment, one strategy is to contrast the admission and mid-stay scores to determine whether the patient worsened, improved, or remained the same. For this purpose, admission scores of 0 and 1 are equated with no morbidity on the mid-stay review, while a 2 corresponds to morbid, and 3 and 4 correspond to major morbidity. In our analyses, we examined the predictive power of the admission MedisGroups score alone, the mid-stay review score alone, and the admission and mid-stay review scores considered together.

Description of the Data Base

The data base used for this study was created through the Health Care Financing Administration (HCFA) Health Standards and Quality Bureau. The data were collected by the peer review organizations (PROs) in Alabama, Arizona, Indiana, New York, Pennsylvania, Utah, and Wisconsin; the states were chosen to provide a geographically representative sample. Cases were randomly selected using administrative data on all Medicare beneficiaries, including disabled individuals younger than 65 years of age, admitted between January 1985 and May 1986. Selection was predicated on the presence of specified diagnostic and procedural codes. Final case selection was dictated by the PROs based on such factors as the availability of medical records. The analytic data base included 20,985 cases from 833 acute care hospitals across the seven states. The number of cases per hospital ranged from one to 480, with a median of 13 cases (mean = 25, standard deviation = 39) per hospital.

Conditions were selected based on their volume, relevance to Medicare beneficiaries (such as the elderly), and complexity of clinical issues raised in their care. The six study conditions were pneumonia or chronic obstructive pulmonary disease (COPD); coronary artery revascularization (including coronary artery bypass graft surgery and percutaneous transluminal coronary angioplasty); acute myocardial infarction; congestive heart failure; cholecystectomy; and prostate surgery (including transurethral and open prostatectomies). Cases in three conditions (pneumonia/COPD, acute myocardial infarction, and congestive heart failure) were selected using the principal diagnosis, while selection in the remaining conditions was based on procedure codes. Because of this selection approach, it was possible for cases chosen by diagnosis codes to have had a major procedure as well.

MedisGroups information and other selected clinical variables were collected by PRO abstractors trained by MediQual Systems, Inc. MediQual requires that reviewers achieve 95 percent exact agreement for items abstracted from the medical record; however, we did not independently confirm the reliability of the data. Data entry was performed by the PROs, then transmitted to MediQual Systems, Inc., for creation of a data tape for HCFA. The tape included admission and mid-stay scores, as well as the values for all KCFs.

Definition of Mortality

The data included date of death information from the Social Security administrative files, making it possible to identify deaths that occurred following discharge. The analyses focused on death within 30 days of admission, regardless of whether the death occurred in hospital. The 30-day figure is currently used by HCFA in producing its annual hospital mortality statistics, and it minimizes concerns about how differences in discharge practices across institutions and regions influence in-hospital death information.[2]

Statistical Methods

All analyses were performed within each of the six conditions. Each of the three models used the MedisGroups scores as categories indicated by dummy variables: admission score, five categories (0, 1, 2, 3, and 4); mid-stay review score, four categories (no mid-stay review, no morbidity, morbid, and major morbidity); and the joint categorization by both the admission and mid-stay scores, 20 categories (the five admission score classes crossed with the four mid-stay score classes). Given that each of our models contained only dummy variables as predictors, ordinary least squares and logistic regression methods produced exactly the same predicted probability of dying for every case (i.e., the same fraction dying among all cases in the same category). We looked at both the R-squared values and the C statistic from a logistic regression[12,13] to examine the ability of the three models based on MedisGroups scores to explain differences in death within 30 days of admission. However, only the R-squared values are reported here since no substantive differences emerged from looking at the C statistic.[10]

Results

Death Rates across MedisGroups Scores

The fraction of patients dying within 30 days of admission varied by condition,

TABLE 2—Number of Cases by Mid-Stay Review Score and Percent Dying within 30 Days by Condition

Condition	No Mid-Stay Review	Mid-Stay Review Score			Total
		No Morbidity	Morbid	Major Morbidity	
	Number of cases (percent dying within 30 days)				
Pneumonia or COPD	214 (38.3)	2,327 (5.5)	655 (13.4)	356 (43.8)	3,552 (12.8)
Coronary revascularization	102 (31.4)	1,837 (2.5)	988 (3.8)	440 (18.0)	3,367 (5.8)
Acute myocardial infarction	412 (72.8)	1,386 (6.4)	1,044 (12.8)	470 (56.2)	3,312 (23.8)
Congestive heart failure	249 (34.1)	1,735 (6.8)	1,127 (12.8)	413 (43.3)	3,524 (14.9)
Cholecystectomy	202 (0.0)	2,581 (1.2)	632 (3.6)	174 (14.4)	3,589 (2.2)
Prostate surgery	228 (0.0)	3,217 (0.8)	165 (4.8)	31 (19.4)	3,641 (1.1)

TABLE 3—Variation in Probability of Death within 30 Days of Admission Explained (R-Squared) by Admission Score Alone, by Mid-Stay Score Alone, and by Admission Plus Mid-Stay Score, by Condition

Condition	Admission Score Alone	Mid-Stay Score Alone	Admission Plus Mid-Stay Score
Pneumonia or COPD	0.11	0.15	0.24
Coronary revascularization	0.04	0.08	0.16
Acute myocardial infarction	0.16	0.34	0.41
Congestive heart failure	0.06	0.12	0.18
Cholecystectomy	0.01	0.04	0.05
Prostate surgery	0.01	0.03	0.05

with the surgical conditions having lower death rates (1.1–5.8 percent) than the medical conditions (12.8–23.8 percent), (Table 1). The distribution of cases across admission severity scores also varied by condition. For example, only 6.1 percent of the surgical cases had high admission scores of 3 or 4, compared with 30.4 percent of the medical cases. Within all six conditions, higher admission scores were generally associated with higher death rates.

However, even among cases with identical admission scores, those in conditions with higher overall death rates generally had higher severity-score-specific death rates. For example, among the six conditions, prostate surgery cases had the lowest overall death rate (1.1 percent) while acute myocardial infarction cases had the highest (23.8 percent). Not only did prostatectomy contain far fewer cases with scores of 3 and 4 than did acute myocardial infarction (1.6 percent of cases compared with 33.2 percent), but also, within each of the five admission scores, the rate of dying was less than one-fifth as great for prostatectomy cases as for infarction cases.

The rates of death by mid-stay review score also varied widely across conditions (Table 2), as did the distribution of cases across mid-stay scores. Even with major morbidity, patients in the three surgical conditions were much less likely to die than their medical counterparts. Death rates rose somewhat between no morbidity and morbid, but the rates increased at least threefold between morbid and major morbidity.

Except in cholecystectomy and prostate surgery cases, many of the patients with no mid-stay review died. For instance, 72.8 percent of acute myocardial infarction patients without a mid-stay review died, as did 31.4 percent of coronary revascularization patients. Thus, the reason for the absence of a mid-stay review was often the early death of the patient. Nevertheless, most of the patients who died remained alive long enough to have a mid-stay review. In three conditions, the plurality of patients who died had major morbidity. In cholecystectomy and prostate surgery patients, the 1 percent death rate among the many cases with no morbidity represented more deaths than the over 14 percent death rates among the few cases with major morbidity. In contrast, 33.6 percent of acute myocardial infarction deaths occurred among patients with major morbidity, while 38.1 percent of deaths occurred among patients with no midstay review.

Explaining Variations in Rates of Death

When admission scores were used to predict 30-day mortality (Table 3), the R-squared values ranged from 0.01 in cholecystectomy and prostate surgery cases to 0.16 in acute myocardial infarction cases. In comparison, the mid-stay review score was a more powerful predictor of risk of death, with R-squared values ranging from a low of 0.03 for prostate surgery to a high of 0.34 for acute myocardial infarction. Combining information about the admission and mid-stay scores generally yielded much greater predictive power than did either score alone, producing the highest R-squared value in all conditions.

Given the large fraction of deaths among those without a mid-stay review, we looked at whether the fact of a mid-stay review alone was predictive of death. We regressed 30-day mortality on a dichotomous variable indicating the presence or absence of a mid-stay review. For acute myocardial infarction cases, the R-squared value produced was 0.19. For other conditions, however, the R-squared value was much smaller: 0.04 for pneumonia/COPD and coronary revascularization, 0.02 for congestive heart failure, and 0.00 for cholecystectomy and prostate surgery. Thus, in all conditions, the simple fact of a mid-stay review did not appear to cause the greater predictive power of the mid-stay review categories.

Discussion

The superior performance of the two approaches involving the mid-stay review was not primarily related to the fact that the absence of a mid-stay review identified patients who died quickly, except among acute myocardial infarction patients. Probably more important is the reliance of the mid-stay review score on physiological derangements generally one week into the hospitalization. It seems almost a clinical truism that patients with severe problems this far into the admission would have worse prognoses than patients without such findings. By contrast, however, admission scores include severe physiological derangements that can often be treated effectively with current medical therapeutics. This conjecture is consistent with the relatively weak performance of the admission MedisGroups score as a predictor of mortality.

In general, patients in the three surgical conditions were much less likely to die, despite high admission scores. The reason may be that those acute physiological derangements that present on admission (and are reflected in high admission scores) represent manifestations of the patients' underlying disease, which will be explicitly addressed by the surgery, usually with a successful outcome (especially for cholecystectomy and prostate surgery). In contrast, certain physiological derangements stemming from acute and chronic medical ailments may be less amenable to control and resolution.

This study has several limitations. First, the number of deaths was relatively small, especially for the surgical conditions. Thus, the R-squared values reported here (particularly for the combined admission and mid-stay score model) may be inflated by overfitting. Second, the results cannot be generalized to younger patients, although the case selection strategy included a small number of Medicare beneficiaries younger than 65 years of age. Finally, MedisGroups was developed to reflect events during a hospitalization. By focusing on death within 30 days of admission, we included some events that occurred following discharge. When analyses were repeated using in-hospital death, admission scores and mid-stay scores produced R-squared values virtually identical to those for 30-day mortality, except for acute myocardial infarction cases, for which the R-squared value for the mid-stay score was 0.39 for in-hospital death compared with 0.34 for 30-day mortality. Similarly, for the combination of admission and mid-stay scores the R-squared values were only 0.01 to 0.03 higher for predictions of in-hospital mortality, except in acute myocardial infarction cases, for which the R-squared value was 0.47 for in-hospital death compared with 0.41 for 30-day mortality.

Even if the mid-stay review score were perfectly predictive of mortality, it is not at all clear that it would be an appropriate risk adjuster for widespread quality assessment efforts, such as the federal hospital-level mortality figures. The mid-stay review rating could be affected by shortfalls in the quality of care, the very event that mortality statistics seek to identify. Therefore, adjusting mortality figures by the mid-stay review score could mask instances of quality problems.

For such risk adjustment, the appropriate construct is the admission "severity" score, which performs relatively poorly as a predictor of 30-day mortality in this data base. These findings could be contrasted with those produced by other models of admission severity, such as the Medicare Mortality Predictor System[14,15] and the Acute Physiology and Chronic Health Evaluation,[16-18] but such comparisons would be hampered by differences in the data bases, predictive models, and statistical techniques used in the different studies. No published research yet exists comparing the performance of various severity systems as predictors of 30-day mortality on comparable data bases across conditions.

Although these results must be viewed as preliminary and pertinent only to Medicare beneficiaries, they may have important implications for states and regions that have mandated the collection of severity information for assessing hospital quality, especially if one purpose is to examine severity-adjusted mortality rates. For example, Pennsylvania and Iowa now require hospitals to collect patient severity information using MedisGroups, with one goal being to examine severity-adjusted outcomes. Our findings raise the crucial question: What causes the large amount of variation in death rates unexplained by the MedisGroups admission score? Further study is needed to answer this question so that the MedisGroups information can be used in an appropriate and valid fashion. □

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APPENDIX—Description of MedisGroups

Selection of Key Clinical Findings

Assignment of MedisGroups severity scores derives from the number, type, and weights of the Key Clinical Findings (KCFs) identified for each individual case. More than 250 potential KCFs are considered, using patient history; physical examination; and laboratory, radiology, pathology, and procedure results to indicate clinical abnormalities. All KCFs are used in computing admission scores, but only a subset of KCFs is involved in deriving the mid-stay review score.

The admission review aims to reflect both acute clinical derangements and more chronic conditions. It encompasses the first two calendar days of the hospital stay for medical patients. For surgical patients, the timing of the admission review depends on when the operation takes place. The intent is to capture all aspects of the surgery, including pathological findings, as part of the admission review. The worst values obtained during this period are collected to compute the admission score.

The so-called "mid-stay review" aims to identify acute or subacute "morbidity" following the admission period. Although all KCFs are collected and all physiological derangements are considered, the mid-stay review computation disregards most chronic or prolonged manifestations of disease. For example, abnormal arterial blood gases in a patient with COPD are included, whereas the chest radiograph finding of COPD would be excluded. The MedisGroups algorithm classes the KCFs used for this mid-stay rating as "acute" (e.g., hematocrit, glucose, blood pressure, pulse, coma, and ventricular fibrillation on electrocardiogram) or "persistent" (e.g., ascites on ultrasound, extensive burns, congestive heart failure on chest radiograph, disorientation, and pulmonary embolism on arteriography).

For most medical cases, the mid-stay review encompasses days 3 through 7 of the hospital stay. For most surgical cases, this review covers the five days following surgery. The worst values obtained during the mid-stay period are collected, but information on vital signs and 10 laboratory KCFs must come from the last 48 hours of the period. Patients generally do not receive a mid-stay review if discharged alive during the designated period.

Scoring

Individual KCFs are assigned weights or "groups" from 0 to 3; group 3 findings indicate a serious clinical derangement and imminent risk of organ failure. Examples of group 0 KCFs include history of congestive heart failure, heart murmur on physical examination, colonic polyp on colonoscopy, serum sodium greater than 128 mmol/L, systolic blood pressure greater than or equal to 60 mmHg, and spinal cord compression found on computed tomography. There are very few group 3 findings; they include coma or stupor on neurological examination, arrest with resuscitation, burns covering greater than 25 percent of body surface area, hematocrit less than 20.0 percent, and systolic blood pressure less than 60 mmHg. Although computation of scores disregards patient age, special pediatric ranges are specified for selected continuous KCFs, and a number of KCFs apply only to neonates.

Employing the algorithm shown in Table A, admission severity scores from 0 to 4 are computed based on the KCFs identified and their weights (i.e., group assignments from 0 to 3), regardless of diagnosis. Using the designated "acute" and "persistent" KCFs, the mid-stay review score is also computed independently of diagnosis, and it results in three classes: no morbidity, morbid, and major morbidity (Table B).

TABLE A—Meaning and Computation of Admission Scores

Score	Meaning of Score	Computation of Score
0	No significant findings	All other, such as cases with multiple group 0 KCFs
1	Minimal findings, indicating a low potential for organ failure	One or two group 1 KCFs
2	Either acute findings connoting a short time course with an unclear potential for organ failure, or severe findings with high potential for future organ failure	One group 2 KCF *or* Three or more group 1 KCFs
3	Both acute and severe findings indicating a high potential for imminent organ failure	One group 3 KCF *or* Two or more group 2 KCFs
4	Critical findings indicating the presence of organ failure	Two or more group 3 KCFs

TABLE B—Meaning and Computation of Mid-Stay Score

Mid-Stay Score	Computation of Mid-Stay Score
No morbidity	All other
Morbid	One group 2 acute KCF *or* One or more group 2 persistent KCFs
Major morbidity	One or more group 3 acute KCF *or* More than one group 2 acute KCF *or* One or more group 2 acute KCF *and* one or more group 2 persistent KCF

The Impact of HIV-Related Illness on Employment

ABSTRACT

We used structured telephone interviews to determine the extent of work loss following onset of symptoms, the interval between onset of symptoms and cessation of work, and the risk factors for work loss among 193 persons with symptoms of human immunodeficiency virus (HIV)-related illness attending the AIDS Clinic at the University of California, San Francisco, between October 1, 1988, and September 30, 1989. Estimates of the duration of time between onset of HIV-related symptoms and work loss derive from the life table method of Kaplan and Meier. A Cox proportional hazards model is used to estimate the effect of risk factors on the probability of withdrawing from work in each time interval. Eighty-six percent of the respondents worked prior to onset of the first symptom of HIV-related illness; 40 percent were working at the time of the most recent interview, a mean of 958 days later. The total number of hours worked declined by 59 percent during this time. Kaplan-Meier analysis indicates that 50 percent who worked prior to onset of HIV-related illness stopped working within two years and all had stopped within 10 years after onset of the first symptom. (*Am J Public Health* 1991; 81:79–84)

Edward H. Yelin, PhD, Ruth M. Greenblatt, MD, Harry Hollander, MD, and James R. McMaster, MHSA

Introduction

The economic cost of acquired immunodeficiency syndrome (AIDS) has been the subject of numerous studies.[1-28] Several analysts have noted that the largest component of the cost of human immunodeficiency virus (HIV)-related illness is due to lost work rather than to direct expenditures for medical care,[1-3] reflecting onset of this condition among persons in the first half of their careers. However, to date there have been no direct estimates of the magnitude of such work loss. For example, in the first comprehensive study of work loss due to HIV-related illness, Hardy, *et al*,[1] applied estimates from Rivin, *et al*,[4] to persons in their samples. Rivin, *et al*, in turn, relied on physicians' recall of their patients' work status in the three months prior to a telephone interview or, if the patients had previously died, in the three months prior to death. Scitovsky and Rice[2] assumed that any person alive at the end of a year was too ill to work 60 percent of the time and that anyone who had died during the year was unable to work at all in that year. The Rivin, *et al*, estimates depended on physicians collecting and remembering information about work status; the Scitovsky and Rice estimates depended on assumptions about the productivity of individuals with AIDS or AIDS-related complex (ARC), assumptions that they did not test directly.

The present study is designed to provide more direct information about the impact of HIV-related illness on loss of employment, clearly the largest component of the cost of these conditions.

Methods

We collected data from persons attending the AIDS Clinic at the University of California, San Francisco (UCSF), which provides medical care to persons with all stages of HIV-related disease. We use those data to quantify 1) the percentage of those employed prior to the first onset of symptoms with HIV-related illness and the percentage employed at the time of the most recent interview; 2) the change in the length of the workweek between the first onset of symptoms and the time of the most recent interview; and 3) the interval between the first onset of symptoms and the cessation of work activities. In addition, we describe the medical and demographic characteristics of persons with HIV-related illness as well as the characteristics of their jobs that are associated with the cessation of work.

Sampling

The sampling frame for this study consists of all 293 persons with HIV-related illness attending a randomly chosen session of the UCSF AIDS Clinic between October 1, 1988, and September 30, 1989. These 293 persons were offered enrollment in the clinic setting; completed telephone interviews eventually were conducted with 193 of them (66 percent). Study respondents are interviewed three times annually to provide an update on health status, medical care usage, and employment history. As of December 1989,

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Address correspondence and reprint requests to Edward Yelin, Department of Medicine, 350 Parnassus Avenue, Suite 600, San Francisco, CA 94117, tel: (415) 476-9027. The authors are all with the Department of Medicine, University of California, San Francisco. This paper was presented at 5th International Meeting on AIDS, Montreal, Canada, June 1989. This paper, submitted to the Journal March 6, 1990, was revised and accepted for publication May 31, 1990.

TABLE 1—Characteristics of Study Respondents

Characteristic	All Study Respondents (%) (n = 193)	Study Respondents with a Work History (%) (n = 170)
Age[a]	39 ± 8	38 ± 7
Male	96	97
Ethnicity		
White, non-Hispanic	82	81
Black	9	8
Hispanic	6	7
Other	3	3
Highest grade attained		
High school or some college	31	28
Bachelor of arts (BA) degree	49	50
More than BA degree	21	22
Occupation		
Professional/manager		28
Clerical/sales		28
Crafts/operatives/manual		18
Service worker		26
Income before onset of first symptom (000 1989 $s)[a]	27 ± 7	30 ± 7
Risk group		
Male homosexual	85	84
Transfusion	3	2
Intravenous drug user	12	13
Diagnosis		
Acquired immunodeficiency syndrome	55	57
Aids-related complex	28	28
History		
Kaposi's sarcoma	26	29
Pneumocystis carinii pneumonia	34	34

[a] Mean ± standard deviation.

97 respondents had been due for reinterview. Of these, nine had died, two refused continued participation because of illness, nine were lost to follow-up, and 77 were reinterviewed. Four persons who were reinterviewed subsequently died, as did one individual not yet due for reinterview.

Study Instruments

The principal data collection instrument is a structured, closed-ended telephone survey conducted by a trained interviewer who asks respondents to report whether they experience each of 20 common symptoms of HIV-related illness, (e.g., weight loss, fever, night sweats) and when the onset of each occurred; whether they experience each of 20 common medical conditions associated with HIV-related illness (e.g., Kaposi's sarcoma, *Pneumocystis carinii* pneumonia) and when the onset occurred; whether they have been told by a physician that they meet the standard criteria for AIDS or ARC, and when they were told so. The respondents provide a complete record of their current medications and those they have taken in the past for HIV-related illness. Following standard protocols for studies measuring labor force participation,[29] the respondents also provide an employment history for the time at which they noticed the first symptoms of HIV-related illness and at the time of the interview. In the employment history, respondents report whether they were employed, looking for work, out of the labor force, or disabled when they first noticed symptoms of HIV-related illness. If they were working, they report hours worked per week, occupation, and industry, and answers to a checklist of working conditions. They then provide the identical information for the day of the interview. Respondents who stopped working in the interim report the approximate date on which they left employment.

Data Analysis

The goal of the analysis is to describe the employment consequences of HIV-related illness among study respondents. We measure the impact of such illness in terms of actual work status—whether the individal is employed for pay in the labor market and, if so, the hours of employment. Thus, we do not measure the impact of HIV-related illness on those who have lost function but are not gainfully employed.

Estimates of the duration of time between onset of symptoms (or diagnoses) and cessation of work activities derive from the life table method developed by Kaplan and Meier,[30] with the censored cases being those who continued to work until the date of the most recent interview or who have been lost to follow-up. A Cox proportional hazards model[31] is used to estimate the effect of the respondents' demographic and medical characteristics, as well as of their job characteristics on their probability of withdrawing from work in each time interval.

The estimates of time to cessation of work activities reported here may be conservative. First, in attempting to renew contact with some of the respondents, we found that many have moved out of the area so that they could be cared for by parents or siblings and have, therefore, left work; such persons are recorded as being lost to follow-up. Second, of the 14 individuals who died subsequent to the first interview, 13 worked prior to onset of HIV-related symptoms. Of these, 6 stopped working by the last time we interviewed them. For the 7 others, for whom we could not establish the actual date on which work activities ceased, we arbitrarily assigned the date of data analysis (September 1, 1989) for the cessation of work activities. This strategy inflates the estimates of time between onset of symptoms and work loss.

Results

Table 1 shows the demographic characteristics of all study respondents and of those with a work history. Almost all are males, most are non-Hispanic Whites, and more than two-thirds finished college. Although a majority worked in white-collar occupations prior to the onset of HIV-related illness, sizable minorities were employed in service work or manual labor.

Table 2 summarizes the symptomatic history of the study respondents. Ninety-six percent reported having at least one potential symptom of HIV-related illness. On average, the symptomatic respondents experienced their first symptom 958 days prior to the most recent inteview, although the responses ranged from 0 days to more

TABLE 2—Symptomatic History of Study Respondents

Symptom	Percent with Symptom (n = 193)	Duration with Symptom, in Days (Mean, ± Standard Deviation, Range)
One or more symptoms of any kind[a]	96	958 ± 677, 0–3, 931
One or more neurological symptoms[b]	60	604 ± 521, 0–3, 931
One or more symptoms of dementia[c]	31	525 ± 336, 0–1, 353
One or more physically limiting symptoms[d]	45	831 ± 584, 0–3, 210

[a] Includes night sweats, diarrhea, shortness of breath, chronic cough, weakness, fatigue, confusion, memory loss, vomiting, pain, changes in sensation, loss of vision, skin lesions, oral thrush.

[b] Includes subset of confusion, memory loss, changes in sensation, loss of vision.

[c] Includes subset of confusion, memory loss.

[d] Includes subset of shortness of breath, weakness, fatigue, pain, loss of vision.

TABLE 3—Productivity of Study Respondents Prior to First Symptom and at Time of Most Recent Interview

All Workers	Full-time Workers	Part-time Workers
	Prior to First Symptom	
(*n* = 166, 100%)	(*n* = 146, 88%)	(*n* = 20, 12%)
	Hours/Week: Mean, Total, Percent	
38.4, 6,381, 100%	40.7, 5,938, 93%	22.2, 443, 7%
	At Time of Most Recent Interview	
All Workers (*n* = 77, 100%)	Full-time Workers (*n* = 51, 66%)	Part-time Workers (*n* = 26, 34%)
	Hours/Week: Mean, Total, Percent	
34.0, 2,616, 100%	40.4, 2,061, 79%	21.3, 555, 21%

than 10 years. Sixty percent of the respondents reported one or more neurological symptoms lasting slightly more than 600 days on average; fewer than a third reported symptoms of dementia, and just slightly fewer than half reported a physically limiting symptom.

At the time of the first symptom of HIV-related illness, 166 respondents (86 percent), were working and 27 (14 percent), were not. This breakdown is very similar to the 84 percent labor force participation rate among all 35–44-year-olds in the United States.[32] At the time of the most recent interview, however, only 77 respondents (40 percent) were working. This figure represents a decrease of 54 percent in the labor force participation rate of the study respondents. Seventy-three persons, or only 44 percent of those who were employed prior to noticing the first symptom of HIV-related illness, were still working at the time of interview. In contrast, 93 individuals, or 56 percent of those employed prior to the first symptom, were no longer working.

Table 3 summarizes the productivity dynamics among the study respondents. The 166 persons who worked prior to the first symptom of HIV-related illness averaged slightly more than 38 hours a week, or 6,381 hours total; 88 percent of them were employed full time. The full-time workers accumulated 5,938 hours per week, or 93 percent of all hours worked. The 77 persons who were working at the time of the most recent interview averaged 34 hours work per week. Their total work time, 2,616 hours, represents a decrease of 59 percent in productivity, with a much larger share of the total hours being accumulated by part-time workers. In fact, the number of part-time workers increased absolutely (from 20 to 26) and relatively (from 12 percent to 34 percent), as did the number and relative share of part-time hours (from 443 [7 percent] prior to the first symptom to 555 [21 percent] at time of interview).

Table 4 apportions the productivity gains and losses of the study respondents by their work histories over the average of 958 days between the first notice of symptoms and the most recent interview. Four respondents began working after the onset of the first symptom of HIV-related illness and another three increased their hours. Forty-nine persons (29 percent of respondents) did not change their hours at all between first onset of a symptom and the most recent interview. Slightly more than 12 percent continued to work, but with reduced hours. Those who stopped working altogether constitute the largest group of respondents (55 percent), and their lost productivity, 3,510 hours, accounts for 90 percent of the overall reduction in hours.

Figure 1 presents Kaplan-Meier estimates of the time to cessation of employment among persons working at the first notice of symptoms. Among all study respondents who had one or more symptoms and who worked prior to the onset of HIV-related illness, 10 percent had stopped working within 100 days of their first symptom (95 percent CI-5, 15). Within a year, 28 percent had stopped working (95 percent CI-21, 36); by two years, 48 percent had done so (95 percent CI-40, 56); by four years, 69 percent had left work (95 percent CI-59, 79); and by 10 years, none was still employed. Among persons who would ultimately develop neurological symptoms, 30 percent had stopped working prior to doing so; among those remaining, 19 percent had stopped working in the first days after onset of neurological symptoms (95 percent CI-10, 28). No respondent with neurological symptoms has been followed for longer than 1,886 days, by which point the 84 percent who were employed when they developed the first neurological symptoms had stopped working (95 percent-CI 58–100).

Figure 2 provides estimates of the time between specific HIV-related diagnoses and the cessation of work activities. Fourteen percent of persons with AIDS had stopped working prior to receipt of that diagnosis. Thereafter, 10 percent had stopped working in the first days after diagnosis (95 percent-CI 3–16), 40 percent had stopped working within a year (95 percent-CI 29–51), and 93 percent were no longer employed as of 1,986 days after diagnosis (95 percent-CI 80–100), the longest any respondent with AIDS has been followed. Of those with ARC, 40 percent had stopped working prior to formally receiving that diagnosis. Thereafter, the proportion working fell steadily, but at an initially slower rate than that of persons diagnosed with AIDS. Ultimately, however, the proportion of working persons with ARC reached zero. Work loss oc-

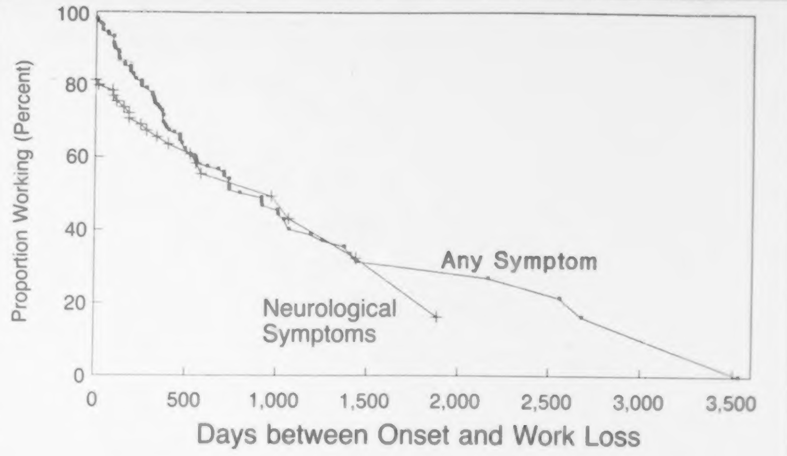

FIGURE 1—Time between Onset of Symptoms and Work Loss

curred more slowly for persons classified as asymptomatic, with 30 percent leaving work by 1,003 days (95 percent-CI 1–59), the longest any respondent formally so classified has been followed.

In a separate analysis, we compared the time between diagnosis of AIDS and cessation of work for persons with Kaposi's sarcoma and those with other conditions that result in the AIDS diagnosis (not shown in Figure 2). This comparison yielded only equivocal evidence that persons with Kaposi's sarcoma are likelier to leave work in each time interval than those with other AIDS-related diagnoses (p = .24 by Wilcoxin test).

In Table 5, we present estimates of the proportional hazard of work loss associated with characteristics of the respondents and of the jobs they held at the time they first noticed symptoms of HIV-related illness. A proportional hazard less (greater) than 1.0 indicates that the person with the risk factor is less (more) likely to stop working in each time interval after onset. Persons with a diagnosis of AIDS stopped working at much higher rates than those who were diagnosed with ARC or who were asymptomatic. Age, a history of male homosexuality or intravenous (IV) drug use, union status, and job tenure did not affect work status. Members of minorities were slightly, though not significantly, more likely to stop working than non-Hispanic whites. The nature of work at the job held at the first onset of symptoms, however, did affect subsequent work status. The proportional hazard of work loss associated with an increment of *each* of the 11 physical requirements of the job was 1.09, indicating that persons with very physically demanding work would have much higher rates of work loss than those in sedentary occupations. Conversely, the proportional hazard of work loss associated with an increment of *each* of the nine measures of discretion in work activities was .93, indicating that persons who controlled the pace and scheduling of work activities were much less likely to leave employment.

Discussion

Both the individuals with HIV-related illness and the society of which they are a part have a strong interest in sustaining the employement of these individuals for as long as possible. Doing so brings such benefits as the sense of well-being they get from continuing in work despite illness, the relatively high income they generate from employment (an average of $30,000 a year prior to the onset of HIV-related illness among the study respondents), and the taxes they pay. The cessation of work on the other hand, initiates one series of entitlements and hastens the day when others will begin. Most persons with HIV-related illness who leave work will apply for unemployment insurance immediately. Those with a diagnosis of AIDS will be eligible for Supplemental Security Income (SSI) as well. If such persons had private disability policies through work, they may begin to collect on them. For those who did not have health insurance through a former em-

TABLE 4—Productivity Gains and Losses

			Hours								
			Prior to First Symptom			Most Recent Interview			Change		
	Persons		Hours/Week			Hours/Week			Hours/Week		
Data Partition	N	%	Mean	Total	%	Mean	Total	%	Mean	Total	% Loss (Gain)
Did not work prior to 1st symptom, worked at interview	4	2	0	0	0	29.0	116	4	29.0	116	(78)
Worked prior to 1st symptom, reduced hours	21	12	42.1	884	14	22.9	480	18	−19.2	−404	10
Worked prior to 1st symptom, same hours	49	29	38.9	1,908	30	38.9	1,908	73	0	0	—
Worked prior to 1st symptom, increased hours	3	2	26.3	79	1	37.3	112	4	11.0	33	(22)
Worked prior to 1st symptom, stopped working before interview	93	55	37.7	3,510	55	0	0	0	−37.7	−3,510	90
Total	170	100	37.5	6,381	100	15.4	2,616	100	−22.1	−3,765	59

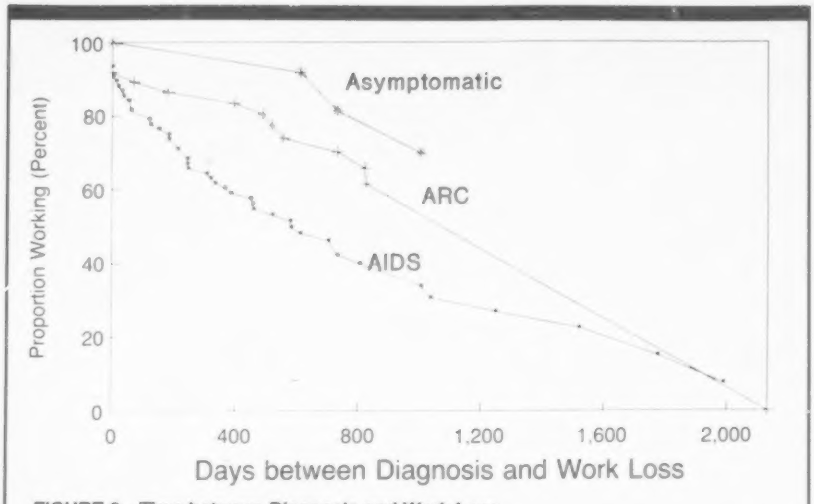

FIGURE 2—Time between Diagnosis and Work Loss

TABLE 5—Impact of Medical and Social Characteristics of Respondents and of Their Jobs on Subsequent Work Status

Characteristic	Proportional Hazard (95% Confidence Interval)	p-value[a]
Age/year	0.99 (0.96, 1.03)	.94
Minority race	1.71 (0.94, 3.13)	.07
Acquired immunodeficiency syndrome	2.73 (1.02, 7.35)	.04
Aids-related complex	1.21 (0.42, 3.52)	.72
Male homosexual	0.61 (0.24, 1.59)	.30
Intravenous drug user	0.30 (0.03, 2.92)	.29
Physical requirements of jobs[b]	1.09 (1.01, 1.17)	.03
Discretion over work[c]	0.93 (0.87, 0.99)	.02
Member of labor union	0.76 (0.40, 1.43)	.38
Tenure on job: years	0.99 (0.99, 1.01)	.99

[a] *p*-value by chi-square test.

[b] Count of physical requirements of jobs, including walking, using stairs, standing or sitting for long periods, stooping, reaching, using eyes for reading or inspection, concentrating or memorizing, using computers, using fingers to grasp, lifting 25 lbs.

[c] Count of discretion over work activities, including pace of work, work procedures, what is produced, time of breaks, sick leave, time of arrival and departure.

ployer, most will be granted Medicaid; those who *did* have health insurance through work, can retain it—at their own expense—only for a year and a half, and then they too will apply for Medicaid. Finally, if they survive the waiting period, they may collect Social Security Disability Insurance (SSDI) and, subsequently, most of those receiving SSDI will become entitled to Medicare.

These potential consequences of work loss following onset of HIV-related illness have been noted before, and there is evidence that the growth in entitlement that follows the cessation of employment has already begun.[4,5,10,18-21,25,26]

There are other compelling reasons as well for understanding the impact of HIV-related illness on employment. As these illnesses become more like other chronic conditions, interventions that maintain function will begin to take precedence over those that merely prolong survival, especially at the cost of function.[33]

There are few data with which to estimate the magnitude of work loss related to HIV-related illness, however. Although the current study is designed to provide direct estimates of such work loss, limitations in the study design may limit the generalizability of these estimates. The study population, although apparently reflective of persons with HIV-related illness in the San Francisco Bay Area,[4,5,34] probably includes a disproportionate number of male homosexuals and persons with a work history compared with the national population of persons with HIV-related illness. Sampling within a clinic setting may also bias the estimates, although after using procedures designed to correct for this problem through weighting,[35-36] we could not detect such a bias and, accordingly, reported the unweighted results above. Finally, the results may be sensitive to the nonparticipation rate.

We found that among all persons with at least one symptom of HIV-related illness, half had stopped working within two years of onset and all had stopped within a decade. Most of the lost productivity is due to total cessation of work rather than to reduced hours. Given that the average age of the study respondents is only 38, most will lose well over half of their careers.

As one might expect, persons who have received a diagnosis of AIDS were at higher risk to stop working in each interval between the first onset of symptoms and the most recent interview than those who have been diagnosed with ARC or who are classified as asymptomatic. Contrary to expectation, risk status—history of male homosexuality or IV drug use—was not associated with work status. Persons who reported that the job they held when they first noticed symptoms of HIV-related illness required more physical exertion and those who reported less discretion over the pace and scheduling of work activities were likelier to withdraw from work in each interval between onset and interview, a finding consistent with studies of work loss after onset of rheumatoid arthritis,[37-38] all forms of musculoskeletal disease,[39] and chronic disease generally.[40] This finding suggests that one way to forestall work loss would be to refer persons with HIV-related illness to less physically demanding work and to jobs in which the employee has discretion over the pace and scheduling of tasks. However, given that the work disability rate eventually reaches unity, any such attempt will have but a temporary effect.

The data reported here reflect the current therapy available at a tertiary care center over the past several years. Advances in treatment could dampen the rate at which work loss occurs among persons with HIV-related illness and, in doing so, prove highly cost-effective. A change in attitudes among employers might alter the willingness of afflicted persons to persevere in the face of illness. The addition of long-term care to the typical workplace-based health insurance policy might remove the incentive to leave work to obtain long-term-care services. In advance of such changes, we find that work loss after onset of HIV-related illness is a pervasive phenomenon. ☐

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Reporting of Occupational Injury and Illness in the Semiconductor Manufacturing Industry

ABSTRACT

In the United States, occupational illness and injury cases meeting specific reporting criteria are recorded on company Occupational Safety and Health Administration (OSHA) 200 logs; case description data are submitted to participating state agencies for coding and entry in the national Supplementary Data System (SDS). We evaluated completeness of reporting (the percentage of reportable cases that were recorded in the company OSHA 200 log) in the semiconductor manufacturing industry by reviewing company health clinic records for 1984 of 10 manufacturing sites of member companies of a national semiconductor manufacturing industry trade association. Of 416 randomly selected work-related cases, 101 met OSHA reporting criteria. Reporting completeness was 60 percent and was lowest for occupational illnesses (44 percent). Case-description data from 150 reported cases were submitted twice to state coding personnel to evaluate coding reliability. Reliability was high (kappa 0.82–0.93) for "nature," "affected body part," "source," and "type" variables. Coding for the SDS appears reliable; reporting completeness may be improved by use of a stepwise approach by company personnel responsible for reporting decisions. (*Am J Public Health* 1991; 81:85–89)

Stephen A. McCurdy, MD, MPH, Marc B. Schenker, MD, MPH, and Steven J. Samuels, PhD

Introduction

Occupational injuries and illnesses represent an important health and economic cost in the United States. The US Department of Labor recorded more than six million cases of reportable occupational injuries and illnesses in 1987, including more than 50 million lost work days.[1] Health and economic consequences for individual workers due to occupational disorders range from trivial to devastating.

Epidemiologic study of occupational disorders has been hindered by the lack of a comprehensive reporting system providing reliable and consistent statistics.[2] The Occupational Safety and Health Act of 1970 attempted to address this problem by mandating the nationwide collection and analysis of statistics on occupational illnesses and injuries meeting specified reporting criteria.[3] The act defines an occupational injury as a disorder that results from an instantaneous exposure (e.g., an impact or fall) occurring in the work environment. Disorders resulting from work exposures that are not instantaneous are considered illnesses. Reporting is mandatory for all cases of occupationally related deaths, all occupational illnesses, and those occupational injuries meeting at least one of four reporting criteria; 1) injuries requiring medical treatment other than first aid, 2) restriction of work or motion, 3) termination or transfer to another job due to the injury, and 4) loss of consciousness.[3]

Each reportable case (and associated lost or restricted work days) is recorded in the company Occupational Safety and Health Administraiton (OSHA) 200 log and classified as an injury or one of the following seven categories of occupational illness: skin diseases, dust diseases of the lungs, toxic respiratory conditions, systemic poisonings, disorders due to physical agents, disorders due to repetitive trauma, and other occupational illnesses. A sample of approximately 280,000 establishments is selected annually to forward a summary of information from the OSHA 200 log to the Bureau of Labor Statistics (BLS) for analysis and subsequent publication.[4] Additional information, including job title and description of the incident and resultant disorder, is recorded on an OSHA 101 form or equivalent state worker's compensation insurance form, which constitutes the Supplementary Data System (SDS). Coding for the SDS is performed by participating state agencies; in California the responsible agency is the California Department of Industrial Relations (CDIR).

In spite of the importance of these data, little research has been conducted to assess the completeness or reliability of reporting for OSHA-reportable conditions. We report here the results of an investigation of reporting of illness and injury in accordance with OSHA reporting criteria among members of the Semiconductor Industry Association (SIA). In addition, two samples of SDS forms were submitted to the CDIR to evaluate the reliability of coding decisions.

Address reprint requests to Stephen A. McCurdy, Division of Occupational and Environmental Medicine, Department of Internal Medicine, Institute for Toxicology and Environmental Health, University of California, Davis, Davis, CA 95616. All the authors are with the Department of Internal Medicine. This paper, submitted to the Journal February 6, 1990, was revised and accepted for publication June 26, 1990.

TABLE 1—Recording by Industry of 101 Reportable Cases

Category[a]	Total Cases n (%)	Re-corded as illness n	Re-corded as injury n	Total Recorded n	Reporting Completeness %
Reportable injury	57 (56)	0	32	32	56
Reportable indeterminate	28 (28)	2	20	22	79
Reportable illness	16 (16)	2	5	7	44
Occupational death	0 (0)	0	0	0	—
Total	101 (100)	4	57	61	60

[a]Determined by authors based on Department of Labor criteria.[3]

Methods

Population

The SIA is a national trade association of semiconductor manufacturing firms (Standard Industrial Classification code 3674). In 1982, the SIA sponsored the development of the Occupational Health System (OHS), a prospective computerized data collection system operating in parallel with the governmental OSHA 200 reporting system. Sixteen companies comprising 37 plant sites and employing more than 96,000 workers, or approximately one-third of semiconductor manufacturing employees in the US in 1984,[4] participated in the OHS. All manufacturing sites participating in the OHS in 1984 agreed to participate in an evaluation of their recording practices.

Company Health Record Review

We evaluated the reporting completeness and reliability of case description data for 1984. Review of cases from the company clinic daily log was conducted at a sample of sites chosen on the basis of a stratified sampling scheme. The 10 sites contributing the most OSHA-reportable cases (median reported cases, 175; median number of employees, 5,951) and a random sample of six of the remaining 27 sites (median reported cases, 24; median number of employees, 920) were selected for site visits.

We examined the daily clinic log for 1984 and selected four work-related cases per month at random from each site. Only cases that the employer had previously determined to be work-related were sampled. When it was available, the abstracter also reviewed the OSHA 101 form (or equivalent state form for the workers' compensation employer's first report), the form for the physician's first report, and the case's personal employee health clinic chart.

If the exposure or event causing the employee's disorder was considered by the abstracter to be instantaneous, the case was classified as an injury. All lacerations, contusions, sprains, strains, and thermal burns were presumed to represent injuries unless the records explicitly indicated that the exposure causing the disorder was not instantaneous. Cases that were likely to have resulted from exposures that were not instantaneous were classified as illnesses. Examples include contact dermatitis due to gloves or creams, acid burns on the hand discovered after the removal of a faulty protective glove, and disorders due to repetitive trauma. Cases in which it could not be determined whether the exposure causing the disorder was instantaneous (e.g., "acid burn on hand") were classified as indeterminate.

We defined the reportability of cases according to BLS guidelines applicable in 1984.[3] The reportability for indeterminate cases was decided according to criteria for occupational injuries. Reporting criteria for injuries and indeterminate cases and work-related death were assumed to be absent unless specifically noted otherwise in the records.

The OSHA 200 form was reviewed to determine if the reportable cases had been recorded. Reporting completeness was calculated as the percentage of cases meeting OSHA reporting criteria that was recorded in the company's OSHA 200 log.

Reliability of SDS Classification

We received 1,117 OSHA 101 or equivalent state workers' compensation forms from a random 50 percent sample of OSHA 200 cases from 1984, as reported previously.[5] Of these, a random sample of 150 forms was submitted to the CDIR without data on the nature, type, source, and body part involved that could identify the person or the company. The initial classifications were stored and the forms were resubmitted without identifying data for repeat classification. The CDIR nosologists received the materials in batches and were not aware that some of the forms had been previously classified. Reliability was calculated as the kappa statistic.[6] Statistical analysis was conducted on a VAX 11/750 computer using the Statistical Analysis System (SAS) software package.[7]

Results

Sixteen sites, representing 66,968 employees, were selected for review. Six sites were not included because they did not have daily clinic logs or supporting records for 1984 suitable for participation in this study, leaving 10 sites representing plants in six states distributed across the US. These 10 sites consisted of seven from the stratum of 10 large case contributors and three from the stratum of six facilities representing the remaining sites.

We reviewed 416 randomly selected work-related cases from the clinic logs. Personal employee health charts were available in 249 (60 percent) of these cases. These records were not available in the remaining cases because of the inability of the companies to locate the records because of loss, transfer, or termination of the employee. There were 101 cases (24 percent) that met reportability criteria for the OSHA 200 form. The median number of reportable cases per site was 8.5 (range: 0–29).

Of the 101 reportable cases, 57 (56 percent) were injuries, 28 (28 percent) were indeterminate cases that met the reporting criteria for injuries, and 16 (16 percent) were occupational illnesses (Table 1). Fourteen (88 percent) of the occupational illness cases represented skin conditions, and the remainder were due to cumulative trauma. Sixty-one (60 percent) of the 101 reportable cases were recorded on the company OSHA 200 form. The completeness of reporting ranged from 13 percent to 100 percent at the nine participating sites with reportable cases. No as-

TABLE 2—Reliability of Classification for 150 Reportable Cases of Occupational Injury or Illness

Classification Variable	Kappa
Nature of injury or illness	0.92
Strain, sprain	0.95
Cut, laceration, puncture	1.00
Chemical burn	0.87
Contusions, crushing injuries, bruise	0.96
Affected body part	0.93
Back	1.00
Finger(s)	0.97
Eye(s)	0.96
Hand (not wrist or fingers)	0.92
Source of injury or illness	0.83
Boxes, crates, cartons, cases	0.97
Hydrofluoric acid	0.96
Bodily motion	0.75
Floor	0.79
Type of injury or illness	0.82
Overexertion in lifting object	0.93
Contact with radiation, caustics, etc. (by skin absorption)	0.94
Fall onto walkway or working surface	0.78
Bodily reaction	0.75

sociation was observed between employee population size or the number of reportable cases at a site and reporting completeness. Because reporting completeness was similar in the two sample strata, the strata were combined for analysis.

Of the 61 reportable cases recorded on the OSHA 200 form, 32 (52 percent) were classified by the abstracter as recordable injuries, 22 (36 percent) were classified as recordable indeterminate cases, and 7 (11 percent) were classified as occupational illnesses (Table 1). Of the 7 illness cases that were reported, 5 (71 percent) were misclassified as occupational injuries. Three of these cases were occupational skin diseases and two were illnesses due to repeated trauma.

For the combined 85 cases of reportable injuries or reportable indeterminate cases, 55 (65 percent) were reportable because of the use of medical therapy beyond first aid and 33 (39 percent) were reportable because of restriction of work motion. Reporting completeness was 78 percent for restriction of work motion cases and 56 percent for medical therapy cases. Reporting completeness was lowest for illnesses. No cases of loss of consciousness or termination or transfer due to an injury and no occupational deaths were observed.

Of the 40 reportable cases that were not recorded on company OSHA 200 logs, 25 (63 percent) were classified by the abstracter as reportable injuries, 9 (23 percent) were occupational illnesses, and 6 (15 percent) were reportable indeterminate cases. The most common reportability criteria among the injury and indeterminate cases were use of medical therapy other than first aid (24 cases) and restriction of work or motion (8 cases). All of the 9 unreported occupational illnesses represented skin disease cases.

Reliability of SDS Classification

The reliability of classification by the CDIR was calculated as the kappa statistic and varied from .82 for type of condition to .93 for affected body part (Table 2). There was little variation in the kappa values between categories within nature, part, type, and source variables.

Discussion

Occupational morbidity imposes an important health and economic burden on individual workers, employers, and society. Primary prevention plays an important role in minimizing this burden. Preventive measures, although generally more cost-effective than treatment, require directed use of resources, and accurate epidemiologic information is needed to guide the application of these resources. The BLS report of occupational injuries and illnesses offers the only nationwide statistics for occupational injuries and illnesses based on standardized reporting criteria. We report the results of our evaluation of reporting completeness within a large and important industry and the reliability of coding for case description data by a state agency participating in the SDS.

A major goal of this study was to evaluate the reporting completeness for conditions that the employer considered work related. Cases that were not recognized as work related would not have been included in the study. True occupational cases may be missed if the employee does not seek medical attention or if the treatment practitioner or employer does not recognize the occupational relationship. Failure to recognize an occupational etiology is particularly likely for illnesses due to chronic exposures and characterized by subtle manifestations or long latencies, such as cancer, neurologic disease, and reproductive disorders. Lack of severity, incomplete medical evaluation, and inadequate recording practices may also lead to underreporting.

In the early 1980s OSHA implemented a site-inspection policy that focused inspections away from establishments that reported low injury rates, raising concern that this policy may promote underreporting. However, examination of these factors and estimation of the true rate of occupational injury and illness would have required active surveillance to capture true cases that would otherwise be undetected and were beyond the scope of this study.

Our findings suggest that the BLS data significantly underestimate incidence of occupational injuries and illnesses. If the 60 percent reporting completeness we observed for conditions recognized by the employer as work related is representative of other industries, the incidence may be nearly 70 percent higher than reported. The underestimation is most significant for occupational illnesses, for which the 44 percent reporting completeness we observed suggests that the incidence may be almost 130 percent higher than reported. The true incidence of occupational conditions is likely to be significantly higher, because the estimates calculated above are based upon observed reporting of conditions recognized by the employer as work related.

Among injury and indeterminate cases, the highest reporting completeness occurred for cases characterized by restriction of work or motion. Injuries and indeterminate cases characterized by the need for medical treatment were less likely to be recognized as reportable. Thus, BLS data are likely to be most valid for conditions such as injuries that directly affect the ability of a worker to perform a job.

The results of this investigation may have been affected by selection bias. The companies studied are members of a large trade association and may not be representative of this industrial sector with respect to reporting performance. Companies not belonging to the trade association may have fewer resources to devote to surveillance and prevention of occupational conditions. Thus, reporting completeness in the group we studied may be higher than for nonmembers. Because the firms participating in this study represent a large fraction of this important industrial sector, we do not believe that any selection bias would have seriously altered our findings and conclusions.

Information bias may also have affected our observed results. Personal employee health charts to supplement data contained in the clinic log were not available for approximately 40 percent of the 416 cases selected for review. Although the clinic logs were generally the most helpful sources of information, personal employee health charts could have contained information indicating a reportable condition. By not reviewing all of these, we may have overestimated reporting completeness. Reporting completeness for occupational illnesses may have been overestimated because indeterminate cases were considered reportable only if they met reporting criteria for injuries. In other words, some indeterminate cases not meeting reporting criteria for injuries may have actually been illness cases but would not have been recognized and classified as reportable in this study. In addition, we presumed that reporting criteria were absent unless they were specifically noted as present; incomplete or incorrect medical record entries could have contributed to erroneous classification with respect to reportability.

Although the influence of these potential biases is difficult to quantify, the data indicate that significant underreporting for occupational disorders, especially illnesses, occurred in the semiconductor manufacturing industry. The increased underreporting for occupational illnesses is significant because the semiconductor manufacturing industry is chemically intensive,[8,9] and illnesses represent a greater proportion of OSHA-reportable cases in this industry than in the manufacturing sector in general.[10] Although the absence of a control industry in this investigation limits comparison with other industries, underreporting has been documented in recent studies of other industries. Eisenberg and McDonald[11] observed an underreporting rate of approximately 20 percent in a pilot investigation of OSHA-reporting practices in 192 manufacturing establishments in Massachusetts and Missouri.

Other investigations have attempted to assess the true rates of occupational conditions. For occupational deaths, estimates of underreporting from 30 to 50 percent have been reported.[12,13] A recent investigation by Blanc and coworkers[14] evaluated the usefulness of occupational illness surveillance using Poison Control Center data. These investigators suggested that the true incidence of occupational illness is three to five times current estimates. Thus, significant underreporting is likely to occur in other manufacturing industries, and a review of underlying causes and possible remedies discussed here may be valuable for industry groups not represented in this investigation.

We also observed misclassification of occupational illnesses as injuries. This phenomenon illustrates important inconsistencies in the OSHA 200 reporting system. Specifically, all occupational illnesses must be reported under the system, regardless of severity, whereas occupational injuries are reported only if they meet specific criteria reflecting severity. Furthermore, classification of illness or injury depends on the exposure leading to the condition. If the exposure was "instantaneous" (a time period defined only by examples of a fall or blow in BLS reporting guidelines), the condition is viewed as an injury. Conditions resulting from all other (i.e., noninstantaneous) exposures are considered illnesses.

Thus, classification of injury and illness in the OSHA system is based on duration of exposure instead of on a biological index such as the nature or severity of the medical condition. Persons unfamiliar with this classification system might use a more biologic and intuitive approach and classify certain illnesses, such as repetitive trauma disorders or chemical burns, as injuries. In addition, the duration of exposure may not be precisely known, and reporting completenes will depend on the interpretation of incomplete or equivocal exposure information by the companies' reporting personnel. However, the majority of underreported or misclassified cases reported here were not equivocal; underreporting appeared due to inconsistent use of published reporting criteria addressing the clinical features of the case.

In spite of the limitations of this study, we have made several valuable observations on which to base specific recommendations. The OSHA data require company personnel at each facility to make case evaluation and classification decisions. Within an industry, the large number of facilities contributing cases introduces the potential of variable reporting practices between facilities, contributing to variability in the quality of the data. In contrast, the coding of clinical data entering the SDS is performed by a small number of trained personnel in a central state office. Our findings indicate that this practice is associated with high reliability of coding for clinical data. It would be impractical to submit all work-related cases to a central office for determination of reportability, and this responsibility is likely to remain at the level of designated personnel at each facility. Reporting practices may be improved and made more uniform by focusing educational measures on these personnel. Education should emphasize recognition of occupational illnesses, for which underreporting was most significant in this study.

We recommend that recording personnel use a formal, stepwise process that forces response to the following pertinent questions to determine reportability of occupational conditions.

- Was the exposure instantaneous (e.g., a fall or a blow)? All cases for which the exposure was not instantaneous are reportable as occupational illnesses. Cases for which the exposure was instantaneous are injuries; the following additional questions must be addressed to evaluate their reportability.
- Was medical treatment other than first aid required, as defined in BLS reporting guidelines?
- Was restriction of work or motion involved that interfered with normal work activities?
- Was there loss of consciousness?
- Was the employee transferred or terminated due to the condition?

If the answer to any of the last four questions is affirmative, the injury is reportable. Regular review of selected cases would help assure consistency and validity of results.

We also recommend that studies evaluating reporting practices be conducted in other industrial sectors. These studies would expand the findings of the investigation reported here and provide a more broadly based picture of the accuracy of national occupational injury and illness data based on the OSHA 200 log. Studies focused on specific occupational conditions should also be undertaken to evaluate the completeness and accuracy of reporting for the OSHA 200 system using independent data sources, such as cancer registries, workers' compensation records, hospital discharge data bases,[15] Poison Control Centers,[14] and occupational health clinics.[16] Accurate surveillance methodology is especially important for occupational illnesses, which are poorly captured by the OSHA 200 system.[13] Finally, case reportability criteria, currently constrained by the Occupational Safety and Health Act, should be reviewed and revised to conform to biological criteria rather than to duration of exposure. □

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Trends in the Content and Use of Oral Contraceptives in the United States, 1964–88

B. Burt Gerstman, DVM, MPH, PhD, Thomas P. Gross, MD, MPH, Dianne L. Kennedy, MPH, Ridgely C. Bennett, MD, MPH, Dianne K. Tomita, MPH, and Bruce V. Stadel, MD, MPH

ABSTRACT

Drug marketing and physician survey data were used to examine trends in the use and hormonal content of oral contraceptives in the United States between 1964 and 1988. Retail prescriptions for oral contraceptives peaked at approximately 68 million in 1973 and have remained between 50 million and 60 million since 1981. Despite this relative consistency in the number of prescriptions, physician "mentions" of oral contraceptives have increased by approximately 75 percent. This increase may reflect closer monitoring of women on oral contraceptives. Use of multiphasic formulations has steadily risen, accounting for 37 percent of the oral contraceptive prescriptions in 1988. Mean estrogen and progestin doses in all types of formulations have steadily declined. A change in the type of estrogen and progestin used in preparations has coincided with this decline in dose. The association between age and use of high-dose formulations seen in the past was no longer evident in 1988. The data demonstrate that oral contraceptive formulations in wide use today differ in hormone content from those of the past, when most of the major studies addressing the risks associated with oral contraceptive use were completed. There is therefore a need to determine the risks and long-term effects associated with these newer formulations. (*Am J Public Health* 1991; 81:90–98)

Introduction

Since oral contraceptives were introduced three decades ago, 56 brands representing 33 unique formulations have been marketed in the United States (see Appendix). Most of these preparations have contained a constant amount of estrogen and progestin (*monophasic* combinations); other types are *sequential*, progestin only ("*minipills*"), and *multiphasic* formulations.

Sequential formulations contained estrogen alone during the first 14 to 16 days of use each cycle followed by an estrogen-progestin combination during the last five to seven days of use. These pills are no longer marketed but were available from 1965 to 1976. Minipills were first introduced in 1972 and have never accounted for more than a fraction of a percent of retail oral contraceptive prescriptions. Multiphasic formulations contain varying estrogen and/or progestin amounts during the cycle and were first introduced in 1982 in an attempt to emulate hormone fluctuations of the menstrual cycle and to lower overall hormone doses.

In considering the beneficial and adverse effects associated with specific oral contraceptive formulations, one must consider the types and doses of estrogen and progestin they contain. Two types of estrogen and nine synthetic progestins have been used in oral contraceptives. The two estrogens, ethinyl estradiol and mestranol, have identical biological effects. Some studies, mostly animal, suggest that on a weight-by-weight basis, ethinyl estradiol is more potent than mestranol in suppressing ovulation, since mestranol must first be converted in the body to ethinyl estradiol.[1–3] In the human body, however, differences in potency between ethinyl estradiol and mestranol do not appear to be important.[4] Progestins contained in oral contraceptives differ by orders of magnitude in potency and may demonstrate varying degrees of andro-genic, estrogenic, antiestrogenic, and other metabolic effects.[5–7]

This paper reports changes in the use of oral contraceptive formulations in the United States from 1964 to 1988. It also examines the relationship between formulation type and the age of users in the period 1980–88. An understanding of these trends is important to better interpret results of previous studies and to plan future studies that seek to clarify the relationship between oral contraceptive use and its potential to cause adverse as well as beneficial effects.

Methods

Sources of Data

Data from two pharmaceutical marketing research data bases produced by IMS America, Ltd. (Plymouth Meeting, PA), were used to examine trends in oral contraceptive use and prescribing. The National Prescription Audit (NPA) was used to provide information on the number or oral contraceptive prescriptions

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Address reprint requests to B. Burt Gerstman, DVM, MPH, PhD, Office of Epidemiology and Biostatistics, HFD-733, Food and Drug Administration, Rockville, MD 20857. The rest of the authors are also with that office of the FDA except Dr. Bennett, who is with the Division of Endocrine and Metabolic Drug Products. This paper, submitted to the Journal January 22, 1990, was revised and accepted for publication May 30, 1990.

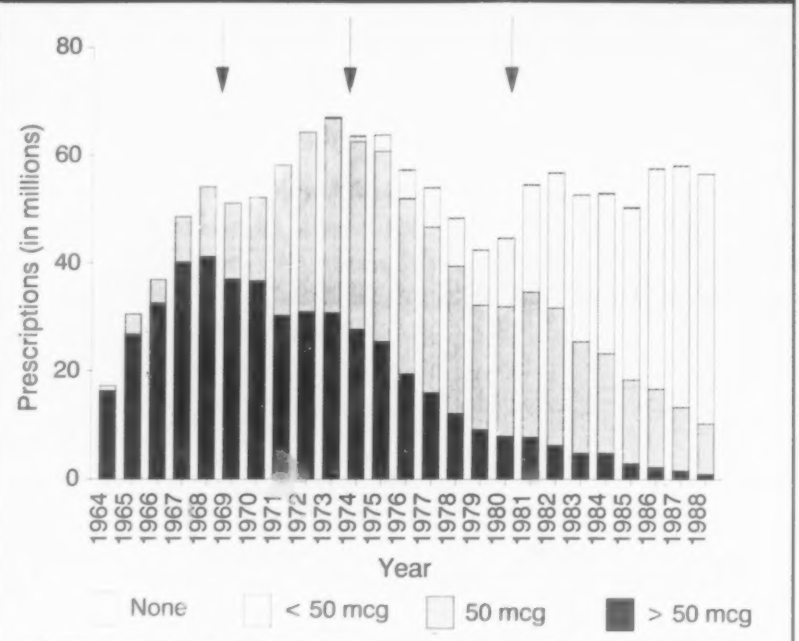

FIGURE 1—Number of Retail Oral Contraceptive Prescriptions by Estrogen Dose, 1964–1988, United States

Arrows indicate points at which changes in data collection methods occurred.

dispensed by retail pharmacies in the contiguous United States. Age-related trends in utilization were based on the National Disease and Therapeutic Index (NDTI) for the years 1980, 1984, and 1988.

The NPA derives its information from chain and independently owned retail pharmacies in the contiguous United States. Other dispensing outlets such as mail order pharmacies, supermarket pharmacies, and clinics are not included. IMS America estimates that retail pharmacies fill approximately 90 percent of prescriptions in the United States. This proportion, however, may not apply to oral contraceptive prescriptions, because approximately one in four women who use reversible forms of contraception rely on publicly funded clinics for care.[8] We were therefore unable to estimate the percentage of oral contraceptive prescriptions that is covered by this study. Nonetheless, previous studies have suggested that trends in oral contraceptive use reported by NPA accurately reflect trends in the general population.[9,10]

From 1981 to 1987, IMS audited all new and refilled prescriptions dispensed at 1,200 computerized pharmacies. This sample was expanded to 2,500 sites in 1987. From 1974 to 1980, data were derived from a representative sample of 800 pharmacies, each of which was audited for two days per month. From 1970 to 1974, 400 pharmacies were audited for four days per month. Prior to 1970, 200 pharmacies reported every fifth prescription filled during the first and third week of each month. Although these changes in methodology may create artifacts at points of change, trends within intervals are unaffected.

Information on the NPA sampling frame is updated annually. Sampled pharmacies represent the universe of retail pharmacies in terms of ownership type, size, and geographic region. Using sampling fractions based on these factors, the number of retail prescriptions dispensed nationwide was estimated.

The NDTI, used to provide information on physician usage patterns of oral contraceptives in relation to age of patients, is based on a sample of 2,130 physicians per quarter recruited from the membership rolls of the American Medical Association and the American Osteopathic Association. The NDTI sample is selected from four geographic regions and 19 specialties, and it is estimated that its sampling universe includes the prescribing practices of 270,827 office-based physicians, but not the prescribing of 51,112 other clinic- and hospital-based physicians. Projection factors derived from the number of physician-days for each of the defined geographic and specialty strata are used to project national estimates.

NDTI-audited physicians are asked to report on each private patient seen or contacted by phone during a specified 48-hour period. Physicians provide information on patient demographics, diagnoses, and treatments for each diagnosis rendered. By convention, the term "mention" is used to denote each drug considered or given during a patient contact. Drugs that are mentioned during a physician-patient contact may reach the patient through prescription, hospital order, provision of samples, or administration or dispensing in the office. It is worthwhile to note that mention of a drug does not necessarily mean that a prescription was issued. Projection factors derived from the number of physician-days in the sample relative to that of the universe for each of the geographically and specialty-defined strata are used to estimate national usage patterns.

Classification of Oral Contraceptives

Formulations were classified by type (monophasic, multiphasic, sequential, minipill), estrogen dose, estrogen contained (mestranol, ethinyl estradiol, none), progestin dose, and progestin contained. Progestins were grouped as follows: norethindrone, ethynodiol diacetate and norethindrone acetate (norethindrone derivative), norethynodrel (oral contraceptives containing norethynodrel were no longer marketed for contraception after October 1988) norgestrel, levonorgestrel, and progestins no longer used in current oral contraceptive formulations (dimethisterone, chlormadinone, medroxyprogesterone acetate). Norethindrone, ethynodiol diacetate, and norethindrone acetate were grouped because previous studies have suggested that these progestins demonstrate similar biological activity and are equally potent on a weight-by-weight basis.[11] We evaluated dose trends by examining both the percentages of retail prescriptions that contained more than 50 µg, 50 µg, and less than 50 µg of estrogen and the mean estrogen and progestin doses of oral contraceptive.

Results

As depicted in Figure 1, total prescriptions rose between 1964 and 1968. After a small decline in 1969, the annual number of prescriptions increased, peaking in 1973. From 1974 to 1979, the number of prescriptions declined steadily. The

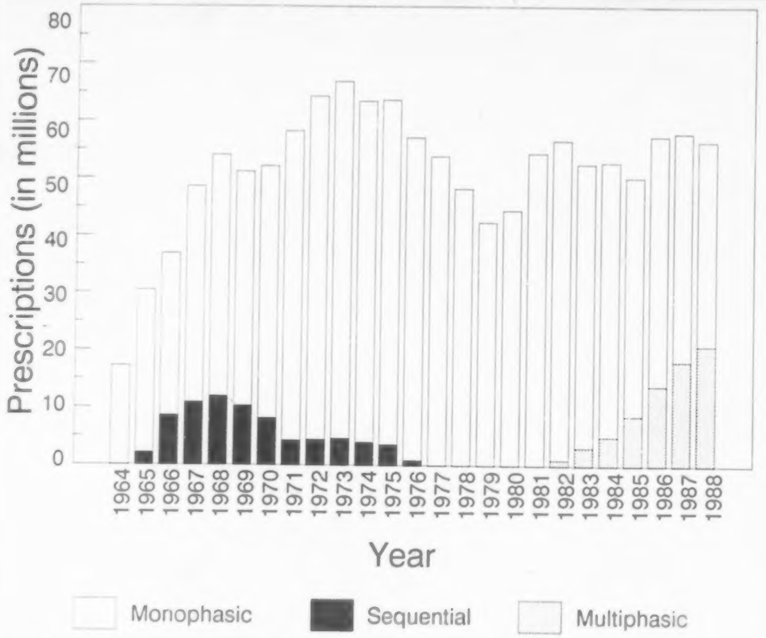

FIGURE 2—Number of Retail Oral Contraceptive Prescriptions by Formulation Type, 1964–1988, United States

TABLE 1—Mean Progestin and Estrogen Dose in Prescriptions for Every Fourth Year between 1964 and 1988, United States

	Progestins				Estrogens	
Year	Norethindrone Group mg	Norgestrel mg	Norethynodrel mg	Levonorgestrel mg	Ethinyl Estradiol μg	Mestranol μg
1964	2.52	—	5.18	—	50[a]	91
1968	1.50	0.50	3.71	—	70	89
1972	1.22	0.50	3.93	—	56	82
1976	1.14	0.46	3.77	—	47	76
1980	1.01	0.41	3.92	—	41	66
1984	0.97	0.37	3.27	0.15	37	62
1988	0.89	0.35	2.73	0.11	35	57

[a] There were only two formulations containing ethinyl estradiol in 1964 (Norlestrin 2.5 mg and Provest), each of which contained 50 μg ethinyl estradiol.

SOURCE: National Prescription Audit, IMS America, Ltd., Plymouth Meeting, PA.

apparent increase in the number of dispensed prescriptions that is seen between 1980 and 1981 could be an artifact of a change in data collection methods that coincided with these dates. Since 1981, the number of prescriptions has remained between 51 million and 58 million.

Prescriptions demonstrate a distinct trend toward use of formulations that contain less than 50 μg of estrogen per tablet (Figure 1). In 1968, fewer than 1 percent of the estimated 54 million retail prescriptions for oral contraceptives were for formulations containing less than 50 μg of estrogen, whereas in 1988 approximately 82 percent of the approximately 57 million retail prescriptions were for formulations with less than 50 μg of estrogen. Retail prescriptions for sequential formulations (no longer marketed) reached their maximum in 1968 (Figure 2). The number of prescriptions for multiphasic formulations steadily increased after they were introduced in 1982, accounting for 37 percent of all retail oral contraceptive prescriptions in 1988.

Mean estrogen and progestin doses have declined over time (Table 1). Since 1981, ethinyl estradiol has become the most common estrogen (Figure 3). Oral contraceptives containing norethindrone derivatives increased market share during the 1960s and have remained fairly constant; they have been included in approximately 65 percent of prescriptions since 1970 (Figure 4). The recent increase in levonorgestrel from less than 1 percent of the market in 1982 to 19 percent in 1988 is attributable to the increased use of two triphasic preparations containing this form of progestin. At the same time, the use of norgestrel declined from 32 percent of the market in 1983 to 16 percent of the market in 1988.

Oral contraceptive mentions by office-based physicians have increased from 1980 to 1988 by 75 percent (13,501 × 10^3 mentions in 1988; 7,716 × 10^3 mentions in 1980: Table 2). The proportion of within-year mentions in the 15- to 24-year-old group has decreased relative to that of the 25- to 34-year-old group. In 1980, 15- to 24-year-old women accounted for 57 percent of oral contraceptive mentions and 25- to 34-year-old women accounted for 37 percent of mentions. In 1988, these figures were 48 percent and 45 percent, respectively. Mentions in the 35- to 44-year-old group accounted for 6 to 7 percent of mentions during this interval.

Mentions for multiphasic formulations increased between 1984 and 1988 (Table 2). In 1984, 15 percent of mentions were for multiphasic formulations. In 1988, 40 percent of mentions were for multiphasics. The largest use of multiphasics occurred in the 15- to 24-year-old age group.

The mean estrogen dose of mentions was directly associated with age of users in 1980 and 1984 (Table 3). The association between age and ethinyl estradiol dose was no longer evident in 1988; the average ethinyl estradiol dose in mentions was 35 μg, irrespective of age. The association between age and mestranol dose has weakened. In 1988, the mean mestranol dose for 15- to 24-year-old women was 56 μg, for 25- to 34-year-old women it was 54 μg, and for 35- to 44-year-old women it was 63 μg.

In 1980 and 1984, 35- to 44-year-old oral contraceptive users were more likely to use mestranol-containing formulations than their younger counterparts. By 1988, only 8 percent of mentions in 35- to 44-year-old women were for mestranol-containing formulations. This figure compares with 15 percent in 25- to 34-year-old women, and 6 percent in 15- to 24-year-old women.

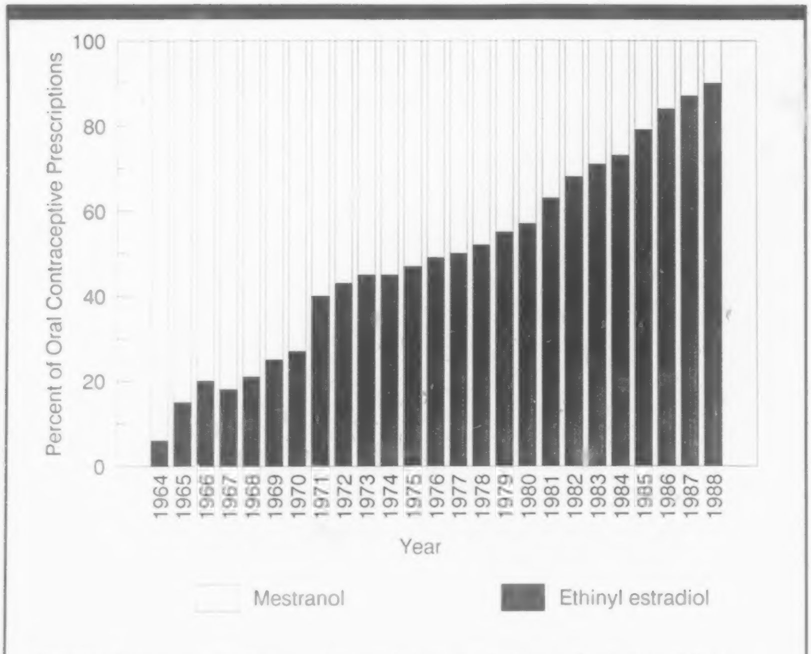

FIGURE 3—Percent of Dispensed Prescriptions by Estrogen Type, United States

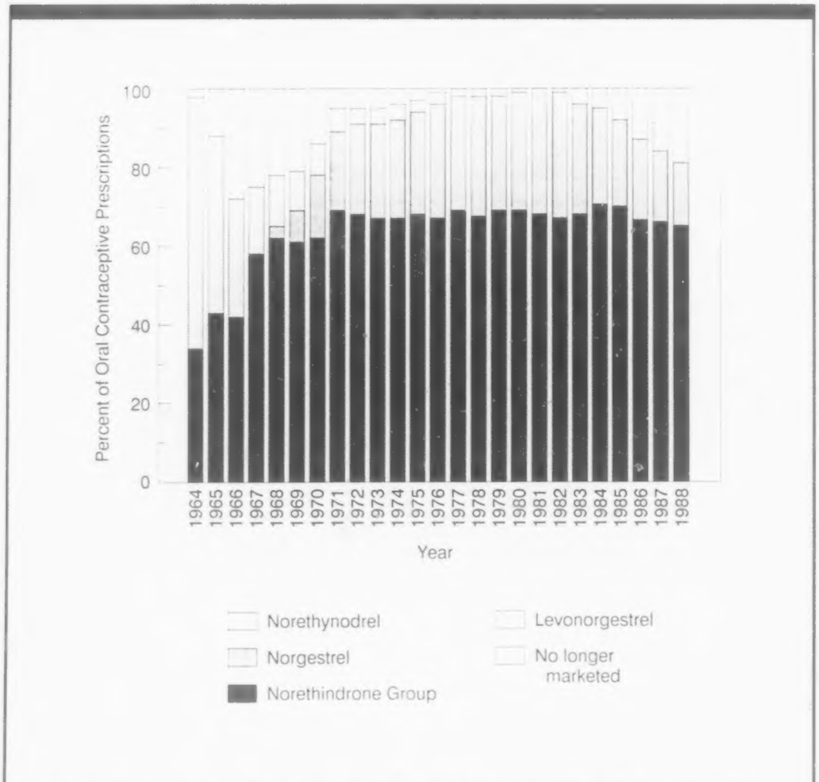

FIGURE 4—Percent of Dispensed Prescriptions by Progestin Type, United States

Discussion

Data presented in this paper pertain to office-based physicians and exclude public clinics as a source of oral contraceptives. In addition, NPA data reflect the number of dispensed retail prescriptions but do not contain information on the number of cycles covered by each prescription. Therefore, NPA data cannot be used to estimate the prevalence of oral contraceptive use in the population. An additional limitation is that there has been no adjustment for change in the number of women of childbearing potential in the population. Despite these limitations, NPA and NDTI data are useful for monitoring large changes in the use of various oral contraceptive formulations over time.

Retail prescription data suggest that overall use of oral contraceptives has been fairly stable since 1981. NPA projections suggest a modest 2.4 percent increase from 1982 to 1987, compared with an 18.5 percent increase in the prevalence of oral contraceptive use reported by Forrest and Fordyce.[12] (Forrest and Fordyce report the percent distribution of women using oral contraceptives increased from 27 percent in 1982 to 32 percent in 1987.) However, neither of these estimates allows for sampling variation, and if we were to estimate the percent change in NPA from 1983 to 1987, a 10.3 percent increase would be noted. This inference, of course, is based on the assumption that both estimates are representative and unbiased and that the average duration of NPA prescription has remained constant over the time interval. However, direct comparisons between NPA projections and those offered by Forrest and Fordyce may not be justifiable, because the former refer to the number of retail prescriptions and the latter are based on self-reported use. Thus, these data are not fully comparable.

Although the number of dispensed retail prescriptions has been relatively stable, the number of NDTI mentions has increased by 75 percent. This apparent inconsistency may be the result of changes in survey methodology, or survey inaccuracy, or it may reflect a change in physician care of patients taking oral contraceptives. We speculate that the observed increase in mentions with no large increase in the number of dispensed prescriptions could be indicative of closer monitoring of women on oral contraceptives.[13]

The impact of the steady increase in the use of multiphasic formulations is unclear because the risks and benefits associated with them compared with other low-dose formulations are largely unknown. Although one article reported a series of functional ovarian cysts in women using multiphasic oral contraceptives,[14] previous studies have demonstrated a reduction in the incidence of

TABLE 2—Number of Mentions of Oral Contraceptives (in Thousands) with Percent of Mentions That Were for Multiphasic Formulations, by Age for 1980, 1984, and 1988, United States

	1980		1984		1988	
Age (yrs)	Total (%)[a]	Multiphasics (%)[b]	Total (%)[a]	Multiphasics (%)[b]	Total (%)[a]	Multiphasics (%)[b]
15–24	4,412 (57)	NA	5,878 (52)	1,009 (17)	6,475 (48)	3,074 (48)
25–34	2,859 (37)	NA	4,741 (42)	636 (13)	6,147 (45)	2,039 (33)
35–44	445 (6)	NA	577 (5)	43 (7)	869 (7)	272 (31)
15–44	7,716 (100)	NA	11,196 (100)	1,688 (15)	13,501 (100)	5,385 (40)

[a] Percent of total oral contraceptive mentions.
[b] Percent of within-age mentions that were for multiphasic formulations.
NA = Not Available

SOURCE: NDTI, IMS America, Ltd., Plymouth Meeting, PA.

TABLE 3—Mean Estrogen Dose of Mentions by Estrogen Type and Age of Patients, United States, 1980, 1984, and 1988

	1980			1984			1988		
Age (yrs)	Mean EE Dose μg	Mean ME Dose μg	% of Mentions Containing ME	Mean EE Dose μg	Mean ME Dose μg	% of Mentions Containing ME	Mean EE Dose μg	Mean ME Dose μm	% of Mentions Containing ME
15–24	39	61	31	35	58	17	35	56	6
25–34	40	64	41	37	61	23	35	54	15
35–44	43	69	53	41	69	22	35	63	8
15–44	40	63	46	36	60	80	35	55	10

EE = ethinyl estradiol;
ME = mestranol

SOURCE: NDTI, IMS America, Ltd., Plymouth Meeting, PA.

functional ovarian cysts in women using high- and intermediate-dose formulations.[15,16] No study to date has measured the risk of functional ovarian cyst formation in users of multiphasic formulations relative to nonusers or to women using other low-dose formulations. Gaspard and Lepot have reported no significant difference in follicular development between users of a low-dose monophasic formulation and its triphasic analog.[17]

Progestin doses have been declining over time. However, all progestins are not equipotent on a milligram-per-milligram basis. For example, previous studies suggest that norgestrel is 5 to 10 times as potent as norethindrone, and levonorgestrel is 10 to 20 times as potent as norethindrone.[10] Using these factors to adjust mean dose, today's norgestrel and levonorgestrel doses are comparable, in terms of potency, to norethindrone doses of the 1960s and 1970s (Table 1). In addition, norgestrel and levonorgestrel may demonstrate more androgenic activity than norethindrone derivatives.[18] Changes in oral contraceptive progestins may have effects on ratios of high- to low-density lipoproteins,[19] glucose tolerance, blood pressure,[20] and the risks of vascular disease,[21–23] benign breast disease,[24] and perhaps even breast cancer.[25]

A decline in the mean estrogen dose contained in formulations was observed. Because the intravascular coagulation and the activity of several clotting factors appear to be estrogen dose dependent,[26–28] this decline has probably resulted in a decline in the incidence of adverse cardiovascular events associated with oral contraceptive use.[29–31]

The correlation between age and high-dose estrogen oral contraceptives seen in the past[32] has almost disappeared. For the most part, women of older reproductive age who choose to use oral contraceptives have taken advantage of the risk reductions associated with low-dose estrogen formulations.

Our data confirm that the oral contraceptives of today contain less hormone than those of the past, when most of the major studies that addressed the risks associated with oral contraceptive uses were completed. Some of these previously defined risks may not necessarily be associated with today's formulations. Additional risks, if any, are not fully known. There is therefore a need to determine the acute and long-term effects associated with these newer formulations. ☐

Acknowledgments

The views in this paper are those of the authors and not necessarily those of the Food and Drug Administration.

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APPENDIX— Hormonal Content of Oral Contraceptives Marketed in the United States, 1964–88

Brand	Years	Progestin	mg^a	Estrogen	μg^a
Monophasics					
Brevicon	1975–88	NEI	0.5	EE	35
Demulen 1/35	1982–88	ETH	1.0	EE	35
Demulen 1/50	1970–88	ETH	1.0	EE	50
Enovid 10 MG	1964–87	NEY	10.0	ME	150
Enovid 5 MG	1964–88	NEY	5.0	ME	75
Enovid-E	1964–88	NEY	2.5	ME	100
Genora 0.5/35	1988–88	NEI	0.5	EE	35
Genora 1/35	1987–88	NEI	1.0	EE	35
Genora 1/50	1987–88	NEI	1.0	ME	50
Gynex 0.5/35	1988	NEI	0.5	EE	35
Gynex 1/35	1987–88	NEI	1.0	EE	35
Levlen	1986–88	LEV	0.15	EE	30
Loestrin 1.5/30	1973–88	NEA	1.5	EE	30
Loestrin 1/20	1973–88	NEA	1.0	EE	20
Lo/Ovral	1975–88	NOR	0.3	EE	30
Modicon	1975–88	NEI	0.5	EE	35
Nelova 0.5/35	1988–88	NEI	0.5	EE	35
Nelova 1/35	1987–88	NEI	1.0	EE	35
Nelova 1/50	1988–88	NEI	1.0	EE	50
Nordette	1982–88	LEV	0.15	EE	30
Norethin 1-35E	1988–88	NEI	1.0	EE	35
Norethin 1-50E	1988–88	NEI	1.0	EE	50
Norinyl 1/35	1980–88	NEI	1.0	EE	35
Norinyl 1/50	1967–88	NEI	1.0	ME	50
Norinyl 1/80	1969–88	NEI	1.0	ME	80
Norinyl 2 mg	1964–88	NEI	2.0	ME	100
Norlestrin 1 mg	1967–88	NEA	1.0	EE	50
Norlestrin 2.5 mg	1964–88	NEA	2.5	EE	50
N.E.E. 1/35	1988–88	NEI	1.0	EE	35

(Continued)

APPENDIX—Continued

Brand	Years	Progestin	mg[a]	Estrogen	µg[a]
Ortho-Novum 10 mg	1964–80	NEI	10.0	ME	60
Ortho-Novum 1/35	1980–88	NEI	1.0	EE	35
Ortho-Novum 1/50	1967–88	NEI	1.0	ME	50
Ortho-Novum 1/80	1968–88	NEI	1.0	ME	80
Ortho-Novum 2 mg	1964–88	NEI	2.0	ME	100
Ovcon-35	1976–88	NEI	0.4	EE	35
Ovcon-50	1976–88	NEI	1.0	EE	50
Ovral	1968–88	NOR	0.5	EE	50
Ovulen	1968–88	ETH	1.0	ME	100
Provest	1964–69	MED	10.0	EE	50
Zorane 1/20	1974–77	NEA	1.0	EE	20
Zorane 1.5/30	1974–79	NEA	1.5	EE	30
Zorane 1/50	1973–77	NEA	1.0	EE	50
Sequential Formulations					
C-Quens	1965–72	CHL	2.0	ME	80
Noriday	1968–69	NEI	1.0	ME	50
Norquen	1967–76	NEI	2.0	ME	80
Oracon	1965–76	DIM	25.0	EE	100
Ortho-Novum SQ	1967–76	NEI	2.0	ME	80
Progestin Only Formulations					
Micronor	1973–88	NEI	0.35	—	0
Nor-Q.D.	1973–88	NEI	0.35	—	0
Ovrette	1974–86	NOR	0.075	—	0
Multiphasic Formulations					
Nelova 10/11	1988	NEI	0.76	EE	35
Ortho-Novum 10/11	1982–88	NEI	0.76	EE	35
Ortho-Novum7/7/7	1984–88	NEI	0.75	EE	35
Triphasil	1984–88	LEV	0.09	EE	32
Tri-Levlen	1986–88	LEV	0.09	EE	40
Tri-Norinyl	1984–88	NEI	0.71	EE	35

[a] Average dose is reported for multiphasic formulations.

CHL = chlormadinone;
DIM = dimethisterone;
ETH = ethynodiol diacetate;
LEV = levonorgestrel;
MED = medroxyprogesterone acetate;
NEA = norethindrone acetate;
NEI = norethindrone;
NEY = norethynodrel;
NOR = norgestrel
EE = ethinyl estradiol;
ME = mestranol

The Status of Restrictive Smoking Policies: A Survey of Medical Schools in the United States and Canada

Frances A. Stillman, EdD, and Diane M. Becker, ScD, MPH

ABSTRACT

All schools of medicine in the United States (N = 128) and Canada (N = 16) were surveyed by telephone to determine if they had instituted policies to restrict smoking. Some policy restricting smoking was reported by 80.56 percent of US schools (N = 103) and by 93.8 percent of Canadian schools (N = 15). However, only 52.3 percent of US (N = 67) and 56.3 percent of Canadian medical schools (N = 9) indicated they had formal written policy statements. Only 13 percent of US schools and 19 percent of Canadian schools had banned smoking totally. (*Am J Public Health* 1991; 81:101–102)

Introduction

Systematic investigation of medical schools' smoking policies have not appeared in the public health literature on smoking policy implementation. This study was designed to determine the extent to which medical schools in the United States and Canada have responded to escalating efforts to eliminate smoking from health care training settings.[1-6]

Methods

Designated representatives from all 128 schools of medicine in the United States accredited by the Association of American Medical Colleges and all 16 schools of medicine in Canada with associative membership status[7] were interviewed by telephone by means of a standardized survey instrument. Surveys were directed to the deans' office and other administrative divisions responsible for regulating smoking for faculty, staff, and students. Data were obtained on reported smoking policies and the implementation process. Each institution that reported a definite smoking policy was asked to provide a copy of the written policy for review. Policy documents were reviewed independently by two raters to ascertain the content and level of restrictiveness of the policy. Interrater agreement was 100 percent.

The policies were classified as "common courtesy," designated no-smoking, designated smoking, total ban, or a total ban plus employee-student restrictions (i.e., on hiring practices, admitting smoking students to the medical school, or both).

A policy classified as common courtesy was one in which the prevailing trend was to seek the preference of others in the environment to determine if there were objections to smoking. Generally smoking would be discouraged only if complaints were raised.

A policy classified as designated no-smoking was one in which smoking was prohibited in designated areas (i.e., lecture halls, labs, and classrooms) but permitted in all other areas.

A policy classified as designated smoking prohibited smoking throughout the environment except in varying numbers of indoor designated areas.

A policy classified as a total ban was one in which smoking was prohibited in all of the interior environment; and no areas for smoking were provided.

A policy classified as a total ban plus employee-student restrictions consisted of a total ban on smoking and a decision not to admit students who smoked to the medical school or hire staff who smoked.

Results

The rate of response to the telephone survey was 100 percent. Based on self-reported survey data, 80.5 percent (N = 103) of US schools and 93.8 percent of Canadian schools (N = 15) indicated the presence of a smoking policy. The majority of schools reported policies of the designated no-smoking type.

Only 52 percent (N = 67) of US and 56 percent (N = 9) of Canadian medical schools that reported having a smoking policy acknowledged the existence of written policy documents. Only two US schools reported having the most restrictive type of policy, which both banned smoking and restricted hiring of employees and admittance of students to nonsmokers. Both schools did so based on religious reasons. The majority of US schools (83 percent) and Canadian schools (75 percent) prohibited the sale of cigarettes on their campuses.

Figure 1 summarizes the percentage of schools with a formal written policy statement grouped according to policy type. All schools in the US and Canada that reported having formal written policies sent copies of their written policy statements. Of the 21 percent of US

Address reprint requests to Frances A. Stillman, EdD, the Johns Hopkins Health Resource Program, 1830 Monument Street, 8037, Baltimore, MD 21205. The authors are with the Division of Internal Medicine, Johns Hopkins University School of Medicine. This paper, submitted to the Journal February 14, 1990, was revised and accepted for publication June 5, 1990.

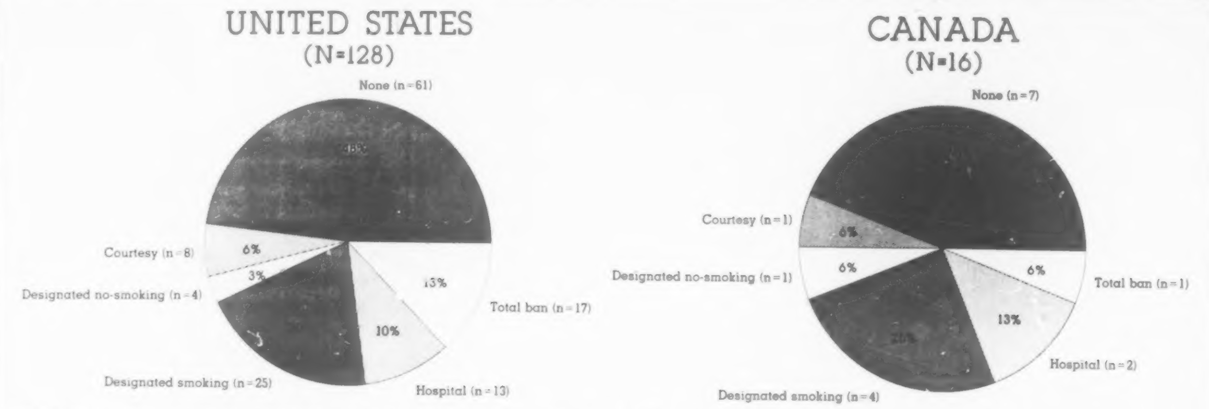

FIGURE 1—Percentage of medical schools with formal smoking policies, by type of policy

schools (N = 27) that reported the existence of a total ban, only 13 percent (N = 17) had formalized the ban into written format. The additional category of "hospital policy" denoted medical schools that had policies that affected only their associated hospitals but not their medical school environment.

Table 1 provides details about the distribution of medical schools with a smoking policy by region based on the written policy statements submitted. The eastern US and the Great Lakes region of Canada had the highest percentage of schools with policies.

While 52 percent of US schools reported providing smoking cessation assistance to students or staff before the policy was implemented, only 36 percent continued to provide some such assistance once restrictions were enforced. The type of assistance provided included formal on-site smoking cessation programs, self-help materials, health screenings, cash incentives for quitting, and referral to community-based smoking cessation programs. In Canada, 53 percent of medical schools provided assistance during the implementation phase of the no-smoking policy.

Discussion

Medical schools with formal policies totally banning or markedly restricting smoking remain in the minority. While the majority of medical schools reported the presence of some smoking policy, barely half of all schools had written policies. The finding that only a relatively low number of schools had made any formal commitment to smoking reduction is surprising given the escalating public interest in smoking and health and reports of successful smoke-free policy implementations in health-care settings.[8,9] By prohibiting smoking, medical schools can make a definitive health statement and act as models of healthy behavior. ☐

TABLE 1—Percentage of US and Canadian Medical Schools with Formal Written Smoking Policies, by Region

Region	Total Number of Schools	School with Smoking Policy N	%
United States	128	67	52.3
West	14	8	57.1
Southwest	11	5	45.5
Midwest	10	5	50.0
Great Lakes	22	11	50.0
South	31	11	35.4
East	36	27	75.0
Other	4	0	0.0
Canada	16	9	56.3
West	5	2	20.0
Great Lakes	5	5	100.0
East	6	2	33.3

Note: U.S. regions consist of AK, CA, CO, ID, MT, NV, OR, UT, WA, and WY in the West; AZ, NM, OK, and TX in the Southwest; IA, IL, MI, MN, OH, and WI in the Great Lakes; IA, KS, MO, ND, NE, and SD in the Midwest; AR, AL, FL, GA, KY, LA, MS, NC, SC, TN, VA, and WV in the South; and CT, DC, DE, MA, MD, ME, NH, NJ, NY, PA, RI, and VT in the East. Other: HI, PR.

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Measurement of Attitudes and Behaviors in Public Health Surveys

François Béland, PhD, Brigitte Maheux, MD, PhD, and Jean Lambert, PhD

ABSTRACT

We divided 900 general practitioners into two groups: one group received a questionnaire measuring both attitudes and behaviors toward preventive aspects of medical care, while the other received two separate questionnaires—one measuring behaviors, and one measuring attitudes. Response rates of the two groups were similar. Respondents in the first group retrieved their answer to the attitudinal items when responding to behavioral items, increasing the correlations between attitudes and behaviors by an average of .147 as compared to the second group; hence the second procedure was preferable. (*Am J Public Health* 1991; 81:103–105)

Introduction

Attitudes and behaviors are salient factors in most of the current public health issues. Most studies measure attitudes and behaviors using a set of items included in a single closed-form questionnaire. Inasmuch as attitudes and behaviors toward the same act are examined, the wording of the items and the response formats should be nearly identical to each other.[1–13] Feldman and Lynch[14] implied that a response to a question is likely to be retrieved as a basis for a subsequent response if they are presented consecutively in one questionnaire. Using similarly worded items to measure attitudes and behaviors compounds the problem. In the present paper we address the question of whether attitudes and behaviors can be measured accurately with a single questionnaire.

Methods

We randomly divided a sample of 900 general practitioners into two groups. One group (A) received a questionnaire in which both their attitudes and behaviors were assessed. The other group (B) received a questionnaire measuring behaviors; on receipt of that completed questionnaire, we mailed them a second one which pertained to their attitudes.

The response rate for group A was 64.3 percent. The response rate for group B was 73.2 percent. Ninety-two percent of the physicians in group B who answered the first questionnaire also completed the second questionnaire.

Twenty-one attitudinal items based on a series of questions (see Appendix) and 21 similarly worded behavioral items were presented to physicians on a six-point rating scale. They measured practices that physicians considered appropriate to perform (attitudes) and those which they claimed to perform (behaviors) in three clinical situations:

- the extent of information that a physician should provide to chronically ill patients about their illness;
- the extent to which a physician should promote non-smoking behavior to patients that consult for reasons other than smoking problems;
- the extent to which a physician should discuss weight with an obese patient consulting for reasons other than obesity.

The items measuring attitudes and behaviors were worded in the same manner.[14] Systematic biases in measuring attitudes and behaviors were assessed by comparing the correlation coefficients between attitudes and behaviors in the two groups of physicians.

The correlation coefficients were transformed into z-scores, using the Fisher Transformation, to test for the difference[15] between the association of attitudes and behaviors in the two groups of respondents. A total of 21 tests were performed for each of the comparisons between two groups. A P-plot procedure was used[16] to estimate the number N of true null hypotheses. The estimate of N was then used to obtain an overall α-level which was less conservative than the usual Bonferonni level. The value of N was 6 in this study, thus the significant level was set at $.05/6 = .008$.

Results

Eight correlations between attitudes and behaviors were different in the two groups at the 0.008 level (Table 1). Pearson's correlation coefficients between attitudes and behaviors were larger by an average of .147 in the group who received one questionnaire than in the group who answered two separate questionnaires.

Some differences in the item variances emerged (data not shown). Because a correlation coefficient is obtained from the division of the covariance with the variances, a small difference in variances can produce large differences in the estimated correlation coefficient. To test whether the differences in the correlations were due to differences in the variances in

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From the Group on Interdisciplinary Research on Health (Béland, Maheux), and the Faculty of Medicine, Université de Montréal (all authors). Address reprint requests to François Béland, PhD, GRIS, Université de Montréal, P.O. Box 6128, Station A, Montreal, Quebec H3C 3J7. This paper, submitted to the Journal June 27, 1989, was revised and accepted for publication May 18, 1990.

TABLE 1—Correlations of Attitudinal and Behavioral Items

Items	Subsamples One Questionnaire	 Two Questionnaires	p-values
Item 1	.502	.352	.039
Item 2	.349	.274	.352
Item 3	.677	.329	.000*
Item 4	.571	.382	.006*
Item 5	.699	.536	.002*
Item 6	.538	.576	.540
Item 7	.594	.482	.028
Item 8	.533	.352	.011
Item 9	.416	.373	.566
Item 10	.366	.317	.546
Item 11	.407	.223	.025
Item 12	.410	.339	.340
Item 13	.693	.468	.000*
Item 14	.470	.386	.248
Item 15	.645	.446	.001*
Item 16	.643	.404	.000*
Item 17	.691	.625	.193
Item 18	.662	.575	.117
Item 19	.655	.436	.000*
Item 20	.642	.365	.000*
Item 21	.472	.319	.041
Mean correlation	.554	.407	

*Different at $\alpha = .05/6 = 0.008$

the items, the 21 variance-covariance matrices for attitudinal and behavioral items of the two groups were compared using the Joreskög procedure.[17] Only four covariances out of 21 were significantly different between the two groups, while 14 variances were significantly different. As a consequence, the differences in the estimated variances had an effect on the estimate of the correlations.

Discussion

The results of our study showed that respondents to a questionnaire, wherein attitudinal items are followed by similarly worded behavioral items, retrieved their answer to the attitudinal items in order to answer questions on behavioral items. Pearsons correlations between attitudes and behaviors were larger by an average of .147 in the group who received one questionnaire than in the group who answered two separate questionnaires.

Some of the differences between the correlations were accounted for by the variances. Thus, fewer covariances than correlations were affected by the data collecting procedures. However, because there were significant differences in the means on the behavioral items between the two groups (table not shown) and because the computation of the covariances depends on the means, the estimates of the covariances were also biased.

Psychometric methods are used in attitudinal and behavioral scale testing. More often than not, researchers use a correlation matrix rather than a covariance matrix when testing for reliability and studying the validity of their scales. Thus, using one or two questionnaires would yield very different reliability coefficients.

The data from this study suggest that when similarly worded items are used to test for both attitudes and behaviors, two questionnaires are required. Nevertheless, the effect of features of questionnaire design should be tested before such a radical conclusion becomes conventional practice. For example, the set of behavioral items was listed immediately after the set of attitudinal items in this study. Had we introduced questions on other topics between the two sets of questions the response bias might have been lowered. Also, it may be possible to use different items to measure both or to use a set of similarly worded and a set of differently worded items. ☐

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APPENDIX—The 21 Attitudinal Items

Items	Questions
Q1	**The general practitioner is often called upon to treat patients suffering from chronic pathologies. When a physician treats a patient suffering from high blood pressure, diabetes or arthritis, for example, should he inform him about:**
	ITEM 1 The medical name of the illness
	ITEM 2 The nature of the illness (e.g. chronicity, diseased organ)
	ITEM 3 The cause of the illness
	ITEM 4 The probable evolution of the illness, if treated
	ITEM 5 The name of the medication prescribed
	ITEM 6 The type of medication prescribed (e.g. diuretics, hypotensor)
	ITEM 7 The respective role of each medication
Q2	**Should the physician do the following things with the types of patients outlined above?**
	ITEM 8 Write a memorandum of the main items of information for the patient
	ITEM 9 Give the patient written documentation on his/her illness
	ITEM 10 Review with the patient at the next visit, the information he/she received from you previously
	ITEM 11 Give the patient the names of resources to contact for additional information on his/her illness (e.g. dietitian, association, etc.)
	ITEM 12 Inform the patient about the good popular books published on his/her health problem
Q3	**Finally, should the physician do the following things according to you?**
	ITEM 13 Try to find out about the patient's social environment
	ITEM 14 Inquire about the worries of the patient vis-à-vis his/her illness
	ITEM 15 Evaluate the support which the patient receives from those who are close to him/her
Q4	**Now let us consider another type of situation frequently encountered in practice. In the case where a young smoker consults for a particular problem unrelated to smoking, do you think the physician should do the following things?**
	ITEM 16 Obtain information on the smoking habits of the patient
	ITEM 17 Check if the patient knows the risks of smoking to his/her health
	ITEM 18 Encourage the patient to stop smoking
Q5	**Finally, in a case where a young obese person consults for a particular problem unrelated to obesity, do you think the physician should do the following things?**
	ITEM 19 Tell the patient what his/her ideal weight should be
	ITEM 20 Offer the patient your help if he/she decides to lose weight
	ITEM 21 Encourage him/her to lose weight

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Determinants of Condom Use among French Heterosexuals with Multiple Partners

Jean-Paul Moatti, PhD, Nathalie Bajos, Jean-Pierre Durbec, PhD, Collette Menard, and Caroline Serrand

ABSTRACT

In September 1988, a sample of French individuals between 18 and 49 years of age, who reported more than one sexual partner in the past six months and who considered themselves heterosexuals (n = 1088), were interviewed at home about risk perception of human immunodeficiency virus (HIV) transmission, sexual behavior, and condom use. Systematic or occasional use of condoms during the previous 12 months was reported by 46.9 percent of respondents. Among condom users, 38.7 percent declared they had never used condoms before the last 12 months, fear of acquired immunodeficiency syndrome (AIDS) rather than contraception being the main motivation of these recent users. Multivariate analysis indicates that voluntary testing for HIV, average or more than average fear of sexually transmitted diseases (STDs), and knowledge of HIV carrier in personal relations are associated with condom use. (*Am J Public Health* 1991; 81:106–109)

Introduction

Since the beginning of the AIDS (acquired immunodeficiency syndrome) pandemic, designing effective prevention has been hindered by a lack of basic information about sexual behavior. While changes in sexual behavior and condom use among homosexual and bisexual men are becoming documented,[1–4] data are still very scarce for the general population.[5] The most recent data on sexual behavior in a large sample of the French population were obtained in a 1969 governmental survey.[6]

France has the highest number of AIDS cases in Europe: 8,025 reported between 1978 and September 30, 1989. A total of 2,176 new cases were diagnosed between October 1, 1988 and September 30, 1989 with 11.5 percent of these cases being due to heterosexual transmission. Although, in France as in the United States, heterosexually acquired AIDS remains focused in demographically distinct population groups (sexual partners of bisexual males, intravenous drug users and individuals of African or Carribean origin),[7,8] prevention of heterosexual HIV (human immunodeficiency virus) spread remains a challenge.

The objective of the present study was to investigate condom use and its determinants in a national sample of French heterosexual population whose sexual behaviors are at risk for HIV transmission. The survey was carried out at a national level in September 1988 and supported by the French Committee for Health Education which was officially in charge of public information campaigns about AIDS.

Methods

The survey data are based on respondents between 18 and 49 years of age, who reported more than one sexual partner in the six months prior to interview and who considered themselves heterosexuals. Individuals were contacted at home. After a short presentation, a written list of various statements corresponding to possible combinations of multiple sexual partners (such as "I am currently married but I had an extramarital sexual relationship during the past six months," etc.) was submitted to the respondents, and they were asked if they recognized having been in at least one of these situations during the past six months. All those who answered positively were included in the survey and were interviewed face to face. The questionnaire included detailed data collection about sociodemographic characteristics and 62 questions about perception of the risk of HIV transmission, sexual behavior, and condom use. The interview lasted between 45 minutes and one hour. Sex of interviewer and respondent were not matched.

A quota sampling procedure for age, sex, occupational status, and size of town was used to select the sample. These quotas were a priori defined on the basis of the observed characteristics of the subgroup of heterosexuals reporting more than one sexual partner during the past six months in a previous survey on AIDS of French general population (December 1987).[9,10] Among the total number of individuals who were contacted at home, 15.3 percent (n = 1,008) and whose characteristics are listed in Table 1 were included in the sample. About half (49.8 percent) were single and a third (34.1 percent) were university graduates.

Univariate statistical analyses and multiple logistic regression analysis were performed with condom use as the dependent variable.

Address reprint requests to Jean-Paul Moatti, PhD, National Institute for Health and Medical Research, INSERM Research Unit on Risk Assessment and Evaluation of Prevention, P.O. Box 48, 92263-Fontenay aux Roses, France. Dr. Durbec is with INSERM Research Unit on Biostatistics and Epidemiology, Marseille; Bajos, Menard, and Serrand were with the Department of Social Studies of the French Committee for Health Education, Minister of Health, Vanves. This paper, submitted to the Journal November 28, 1989, was revised and accepted for publication June 25, 1990.

TABLE 1—Frequency of Condom Use by Characteristics of the Respondents

Characteristics	% of Individuals Having Used Condoms during the Previous 12 Months
Sex	
Male (n = 501)	52.3
Female (n = 507)	41.6
Age (years)	
18–29 (n = 517)	50.7
30–39 (n = 325)	43.7
40–49 (n = 166)	41.6
Marital Status	
Single (n = 502)	52.6
Married or as married (n = 359)	41.2
Separated or widowed (n = 147)	41.5
Size of Town	
Less than 100,000 inhabitants (n = 432)	41.9
More than 100,000 inhabitants (n = 576)	50.7
Level of Education	
University graduates (n = 344)	51.7
College graduates (n = 222)	47.3
Lower level (n = 442)	43.0
Employment	
Executives, professionals and middle management (n = 351)	46.2
Employees and blue-collar workers (n = 423)	47.5
No occupation (n = 234)	47.0
Individual's Perception of Risk of Contracting AIDS	
Greater than average (n = 66)	53.0
Average (n = 235)	52.8
Less than average (n = 495)	46.3
None (n = 212)	40.1
Individual's fear of STDs	
Greater than average (n = 171)	55.0
Average (n = 475)	53.3
Less than average or none (n = 362)	34.8
Past Medical Care for STDs	
Yes (n = 215)	53.5
No (n = 793)	45.1
Testing for HIV	
Voluntary testing for HIV (n = 128)	71.9
HIV testing during systematic screening (n = 127) (blood donation)	44.4
No HIV testing but thought about having a test (n = 322)	47.8
No HIV testing and did not think previously about having one (n = 431)	39.7
Personal Knowledge of a HIV Carrier	
Yes (n = 153)	62.1
No (n = 855)	44.2

Results

Condom Use and Characteristics of Users

In this French sample, 46.9 percent reported having used condoms during the previous 12 months, 10.3 percent always using them. Among these 473 condom users, 38.7 percent declared they had never used condoms before the previous 12 months. These recent users mentioned fear of AIDS as their motivation for condom use in the same proportion as earlier condom users (65.0 percent vs 60.7 percent); however, they reported contraception as a motive for condom use less frequently (31.7 percent vs 62.8 percent).

Condom use is related to several sociodemographic characteristics of respondents as well as risk perception of contracting AIDS and STDs (sexually transmitted diseases) and past medical care for STDs (Table 1). It is of interest that 47.2 percent of respondents who feel being at "average" or "higher than average" risk of contracting AIDS do not yet use condoms (Table 1). Condom use is highest among respondents who voluntarily underwent testing for HIV antibodies and lowest in the non-tested population. It is also more frequent among respondents who personally know an HIV carrier (Table 1).

A minority of respondents (10.2 percent) believe that coïtus interruptus, vaginal douche, or to urinate after sexual intercourse may protect against HIV transmission; but their frequency of condom use is similar to that of the rest of the sample.

Determinants of Condom Use

When a logistic regression model is applied to condom use (either on some or all occasions of sexual intercourse in the last 12 months) and separate regressions are done by sex of respondents, the following variables appear to be related with use: voluntary testing for HIV, average or more than average fear of STDs and knowledge of HIV carrier in personal relations (Table 2).

Perception of Condom Use

Table 3 shows that expressed arguments about individual difficulties to use condoms such as "they are difficult to use" or "they reduce personal interest in sexual intercourse" do not differentiate users from non-users. It is rather criticisms of the social and cultural image of condoms with its potential effect on individual relationships that are more frequently expressed by non-users than users. Non-users also express more doubts about effectiveness of condoms to protect against transmission of HIV and other STDs.

Discussion

Several caveats must be recognized at the outset concerning our data. In the absence of sufficiently large and unbiased samples indicating the real prevalence of various sexual behaviors in the French population, our respondents cannot be considered as fully "representative" of French heterosexuals with multiple sexual partners. Inclusion in the study was based on individuals' willingness to recognize they had more than one sexual partner in the previous six months; so, reliability and validity of answers are difficult to assess (unfortunately, a possible influence of the interviewer's gender on the respondents' answers was not controlled). Our results are very similar to those observed in December 1987 in a representative sample of the population of 18 years of age and over of Paris region: in that survey of the general population, previous testing for HIV, being

TABLE 2—Factors Related to Condom Use in French Heterosexuals with Multiple Partners (logistic regression model)

	Male Respondents (n = 501)		Female Respondents (n = 507)	
Factors present	Odds Ratio	95% CI (*)	Odds Ratio	95% CI (*)
Voluntary testing for HIV	2.46	1.31, 4.59	3.27	1.82, 5.91
Average or more than average fear of STDs	1.44	1.19, 2.55	2.38	1.59, 3.56
Personal knowledge of HIV carrier	1.39	0.82, 2.35	1.95	1.13, 3.35
Age (**)	0.77	0.50, 1.17	1.07	0.70, 1.66
Unmarried	1.34	1.10, 1.85	1.39	0.92, 2.11
University education	1.17	1.00, 1.74	1.48	0.99, 2.20
Prior care for STDs	1.13	0.73, 1.77	0.99	0.61, 1.61

*The confidence interval is not symmetric because it has been calculated on the log odds.
**<30 vs ≥30 years of age.

TABLE 3—Expressed Opinions about Condoms

	Respondents		
% of Respondents Who Agree with the Following Statements about Condoms	Condom Users (%)	Condom Non-Users (%)	Difference (95% CI)
Are difficult to use	18.8	21.7	2.9 (−2.1, 7.9)
Reduce personal interest in sexual intercourse	60.7	64.2	3.5 (−2.5, 9.5)
Induce sexual impotence	5.5	6.2	0.7 (−2.2, 3.6)
Limit one's sexual pleasure	54.8	69.5	14.7 (8.8, 20.6)
Limit sexual pleasure of partners	59.6	62.4	2.8 (−3.2, 8.8)
Kill romanticism in relationship	62.6	76.8	14.2 (8.6, 19.8)
Create doubts about partner	28.1	40.0	11.9 (6.1, 17.7)
Are shameful to buy	7.2	12.3	5.1 (1.5, 8.7)
Should be considered like ordinary consumer goods	70.6	55.9	14.7 (8.8, 20.6)
Are not always effective STD protectors	55.2	66.4	11.2 (5.2, 17.2)

single, level of education, and declaring more than one sexual partner in the past six months were associated with condom use.[9,10]

Another limitation for our data is that those who report more than one sexual partner in the past six months may exhibit a wide variety of sexual behaviors to the risk of HIV transmission. Moreover, sexual-orientation self-reporting (i.e., heterosexual in our study) does not accurately describe an individual's sexual behavior: "heterosexuals," "bisexuals," and "homosexuals" may sexually interact across the boundaries implied by self-reported sexual-orientation.[11,12] Finally, condom use, especially if it is only occasional, is far from being a perfect indicator of protective behavior.

However, the aim of this study was to inform decision-makers in charge of public education campaigns about AIDS rather than to provide in-depth analysis of sexual behaviors. For that purpose, some useful indications and guidelines for general public information about condoms can be derived from our results.

Since 1987, French public health officials have tried to promote condom use through media advertising directed toward the general public (TV ads, newspapers, movies). It is well known that information alone does not cause sufficient change in behaviors, especially those which require significant changes in individuals' habits and self-representations.[13] It has even been argued that public education is particularly ineffective in areas or groups of low HIV prevalence.[14] Our study suggests a less pessimistic view: 18.2 percent of our sample being recent condom users of the last 12 months mainly motivated by fear of AIDS. It indicates that condom use has become relatively more acceptable in some segments of sexually active heterosexual population. Similar results have been found in The Netherlands where promotion of condom use in the general public also began in 1987.[15]

The experience of HIV prevention in the homosexual communities suggests that condom users are less likely than nonusers to believe that condoms "spoil sex" and that interventions should focus on helping people develop strategies and techniques that make condom use sexually satisfying.[16,17] From our results, it seems that other aspects must also be taken into account:

- The importance of fear of STDs as a predictor for condom use suggests that it may be useful to include information about condoms in the general context of STDs rather than to only associate them with prevention of HIV infection. Evaluation of various types of messages will be necessary to further document this point.
- Although expressed doubts about effectiveness of condoms among non-users may be an posteriori rationalization of behavior rather than a real barrier to use, it seems necessary to point out the effectiveness of condoms as a means of prevention and to popularize technical quality controls of the product.
- The most important aspect of general information about condoms may be to change their social image and to promote them as ordinary consumer goods playing a normal role in sexual and affective relationships. That is what French information campaigns have called a strategy for making condoms "more of a social routine." Some evidence observed among IV drug users seems to reach similar conclusions.[18]

French campaigns now issue messages about condoms which do not explicitly associate them with AIDS prevention per se and which emphasize their use as non-conflicting with socially and individually valuable relationships. ☐

Acknowledgments

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Research on the Prevalence and Impact of Drug use in the Workplace: National Institute on Drug Abuse, Research Grants Program, Program Announcement No. PA-90-10, July 1990

The National Institute on Drug Abuse (NIDA) requests proposals from researchers in fields including but not limited to psychology, epidemiology, sociology, business and labor relations, public health, management, and economics, on the prevalence, etiology, and impact of acute, chronic, and "hangover" effects of drug use by the workforce. Specifically, NIDA requests proposals on:

- pacts of drug use on the workplace including absenteeism, accidents, injuries, job turnover, health care costs, and productivity effects;
- job performance assessment;
- simulation and field studies of workplace environments;
- prevalence and etiology of workplace-related drug use;
- prevalence estimation techniques for work-related populations; and
- studies of relationships among worker, job, and organizational characteristics, and patterns of drug use and consequences.

ADAMHA urges the inclusion of women and minorities in study populations. For further information contact: Workplace Performance and Technical Research Branch, Division of Applied Research, National Institute on Drug Abuse, Parklawn Building, Room 9A-53, 5600 Fishers Lane, Rockville, MD 20857, telephone (301) 443-6014.

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Cardiovascular Risk Factor Levels in Ethnic Hawaiians

ABSTRACT

We report a cardiovascular risk factor survey of "native" Hawaiians 20–59 years old (70 percent, or 257), living on the Hawaiian Homestead lands on the island of Molokai, Hawaii. More than 60 percent of both sexes were overweight. Among males, 42 percent were smokers. The percent of the population with systolic blood pressure greater than 140 mm Hg or a diastolic pressure greater than 90 mm Hg or taking hypertensive medications was 14 percent of those ages 20–39 and 36 percent of those ages 40–59. The percent with serum cholesterol ≥ 6.2 mmol/L ranged from 8 percent of those 20–29 years old to 46 percent in those 50–59 years old. Two percent of those ages 20–29 had a history of diabetes, or 2+ or greater glycosuria by dipstick, as did 23 percent of those ages 50–59. The majority of the known diabetics exhibited glycosuria and elevated glycohemoglobin levels, indicating poor control. Hypertension, although usually known to the participant, was frequently uncontrolled. From these data, it appears that among this group of Hawaiians major risk factors for cardiovascular disease were frequent, while at the same time the levels of awareness and/or control for most of these factors were low. (*Am J Public Health* 1991; 81:164–167)

J. David Curb, MD, N. Emmett Aluli, MD, Judith A. Kautz, Helen Petrovitch, MD, Synnove F. Knutsen, MD, Raymond Knutsen, MD, Helen K. O'Conner, and William E. O'Conner, MD

Introduction

In the first years of contact with Westerners, the native Hawaiian population was devastated by epidemics of infectious diseases. Today the life expectancy of native Hawaiians remains the lowest of all the major ethnic groups in Hawaii.[1] The death rate from cardiovascular disease of "pure" Hawaiians is more than twice Hawaii's overall rate.[2] There are, however, few scientifically gathered data on risk factors for cardiovascular disease in native Hawaiians. This study was aimed at providing initial insight into the patterns of cardiovascular risk in a native Hawaiian population.

Methods

A cardiovascular disease risk factor survey of native Hawaiians living on Hawaiian Homestead lands on the largely rural island of Molokai, Hawaii, was conducted in 1985 with strong support from a Hawaiian advisory group. Hawaiian Homestead lands consist of tracts of land allocated by federal law for long-term lease, in individual parcels, to native Hawaiians with at least 50 percent native Hawaiian ancestry. The residents of these areas live primarily in extended family homes. The survey was carried out in conjunction with the INTERSALT project, an international cooperative study on sodium intake and blood pressure.[3] Some data collection instruments and training were common to the two protocols.

There are today relatively few genetically "pure" Hawaiians. Those who identify themselves as Hawaiians have many different lifestyles. Of the major Hawaiian islands, the island of Molokai has the highest proportion of its population who are Hawaiians or part Hawaiians. Because Molokai has experienced fewer of the lifestyle changes associated with mass tourism than other islands, its Hawaiian population, especially that portion living in the Homestead areas, is thought to be more typically Hawaiian. There is widespread unemployment and access to health care is limited.

A list of the individuals, by age and sex, known to be living in the Homestead areas was developed by interviewing longtime residents of the areas, examining hospital and clinic records and voter registration lists, and visiting most homes in the area. Introductory letters, brochures, and newspaper publicity as well as extensive telephone contact and visits to homes by staff and Hawaiian volunteers were used to recruit study participants. The population of individuals 20–59 years of age, excluding pregnant women, was defined as the target population. At the clinic appointment, health, lifestyle (smoking, alcohol use, and physical activity), and de-

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Address reprint requests and communications to J. David Curb, MD, Division of Geriatrics, Depatment of Medicine, John A. Burns School of Medicine, 347 N. Kuakini HPM9, Honolulu, HI 96817. Dr. Curb is with the University of Hawaii John A. Burns School of Medicine, Honolulu, Hawaii, and Na Pu'uwai, Inc., Moloka'i, HI. Aluli and H. K. O'Conner are also with Na Pu'uwai, Inc. Kautz is with the University of Hawaii College of Business Administration, Honolulu, HI. Petrovitch is with the University of Hawaii John A. Burns School of Medicine. Synnove Knutsen and Raymond Knutsen are with the University of Tromso School of Medicine, Tromso, Norway. W. O'Conner is with the University of Oregon Health Science Center, Portland, Oregon. This paper, submitted to the Journal July 25, 1989, was revised and accepted for publication June 1, 1990.

TABLE 1—Number of Individuals Examined, by Age and Sex

Sex	20–59 Years	20–29 Years	30–39 Years	40–49 Years	50–59 Years
Male	123	29	30	37	27
Female	134	27	46	35	26
Total	257	56	76	72	53

mographic questionnaires were administered. Nonfasting blood was drawn for cholesterol, triglyceride, high-density lipoprotein (HDL) cholesterol, and glycosylated hemoglobin measurements. Height, weight, and blood pressure were measured. Spot urine specimens were tested for protein and for glucose by means of dipsticks.

Blood specimens were refrigerated immediately and processed within four hours of collection. Serum specimens were frozen and transported frozen to the laboratory for analysis. Whole blood specimens were transported on ice for glycohemoglobin determinations, which were carried out within three to four days after collection.[4] Standardized Lipid Research Clinic procedures were utilized to determine serum lipid values.[5] Three sitting blood pressures were taken, after five minutes of rest, by individuals certified with the London School of Hygiene audiotape.[3] For each subject's blood pressure value in analyses, an average of the last two blood pressures was used. Hypertension was defined as either taking antihypertensive medications or having a diastolic blood pressure of 90 mm Hg or greater, or a systolic pressure of greater than 160 mm Hg.

Where appropriate, 95 percent confidence intervals (CIs) were calculated for total male and female rates based on the normal approximation to the binomial distribution.[6] These tended to be relatively wide because of the small sample size, and small numbers made the procedure inappropriate for the age-sex cells. Computer analyses of the 1976–80 National Health and Nutrition Examination Survey (NHANES II)[7] and the 1985 National Health Interview Survey (NHIS)[8] were conducted to provide comparative data on a national scale for comparable age groups.

Results

Tables 1 and 2 contain subject frequencies by age and sex and mean values and rates by sex for selected cardiovascular disease risk factors. Among the 367 persons identified as eligible, 257 (70 percent) participated in the study: 61 percent of the 20–29 age group, 72 percent of the 30–49, and 80 percent of the 50–59.

Approximately 65 percent of the participants were 20 percent or more above the average body mass index (weight-height ratio^2) for Caucasians, compared with 27 percent of all US adults ages 20–59.[9] Forty-five percent of the Hawaiians were 40 percent or more overweight by these standards.[9] The mean waist-hip ratio was higher than a 0.80 cut point,[10] indicating upper-body fat predominance, in all sex and age groups among the Hawaiians, but the finding was more marked in males.

Approximately 34 percent (95% CI = 26, 42) of females and 42 percent (95% CI = 32, 51) of males admitted to being current smokers by questionnaire. The highest prevalence of smoking was seen in males ages 20–29 (48 percent). Overall an additional 26 percent of males and 15 percent of females were past smokers.

The prevalence of high blood pressure or taking antihypertensive medications was 24 percent (95 percent CI = 19, 31) in females and 26 percent (95 percent CI = 19, 33) in males. The proportions are similar to those of Whites (Figure 1). Fewer than 7 percent of those meeting the criteria used for blood pressure elevations in these analyses had not been told previously that they had high blood pressure. However, among the 10 percent of the total population taking antihypertensive medications, 81 percent of females and 90 percent of males still had high blood pressures as defined.

The prevalence of individuals in the population with serum cholesterol of 5.2 mmol/L or higher was 50 percent for the males (95% CI = 41, 59) and 45 percent for the females (95% CI = 37, 53). The prevalence of those with serum cholesterol of 6.2 mmol/L or greater ranged from 9 percent in those ages 20 to 29, to 40 percent of those ages 50 to 59. The age-adjusted mean HDL was considerably lower

TABLE 2—Mean Values and Standard Deviations for Selected Cardiovascular Risk Factors in 257 Ethnic Hawaiians, by Age and Sex

	20–59 Years		20–29 Years		30–39 Years		40–49 Years		50–59 Years	
Risk Factors and Sex	Mean	(SD)	Mean	(SD)	Mean	(SD)	Mean	(SD)	Mean	(SD)
Serum cholesterol (mmol/L)										
Male	5.28	(1.24)	4.68	(1.24)	5.25	(1.34)	5.66	(0.96)	5.51	(0.44)
Female	5.25	(1.32)	4.63	(1.01)	5.02	(1.19)	5.30	(1.42)	6.26	(1.22)
Serum HDL-cholesterol (mmol/L)										
Male	1.01	(0.26)	1.06	(0.28)	0.98	(0.25)	0.96	(0.23)	1.03	(0.27)
Female	1.09	(0.24)	1.09	(0.24)	1.06	(0.23)	1.06	(0.24)	1.16	(0.27)
Body mass index[a]										
Male	32	(7.4)	31	(6.0)	32	(4.3)	33	(8.9)	31	(6.7)
Female	31	(8.1)	29	(7.1)	31	(7.9)	32	(8.7)	32	(8.6)
Waist/Hip Ratio										
Male	.95	(0.77)	.91	(0.07)	.93	(0.05)	.98	(0.08)	.97	(0.06)
Female	.84	(0.77)	.81	(0.05)	.83	(0.07)	.86	(0.07)	.86	(0.08)

[a]weight/height ^{2}

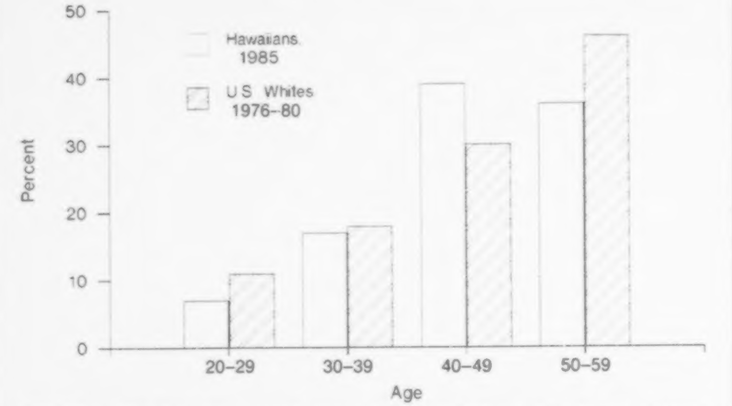

FIGURE 1—High Blood Pressure in Hawaiians vs. US Whites. Criteria of high blood pressure were diastolic blood pressure ≥ 90, or systolic blood pressure ≥ 140, or being on a regimen of antihypertensive medications. Data on US Whites are from the National Health and Nutrition Examination Survey, 1985.

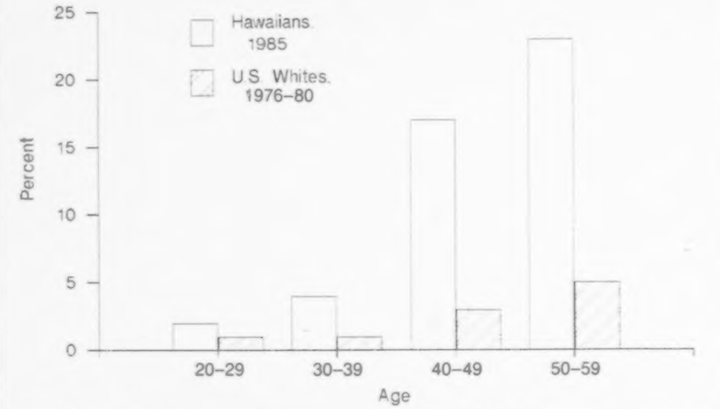

FIGURE 2—Diabetes in Hawaiians vs. US Whites. The percent of Hawaiians who said their doctors had told them they had diabetes or who showed significant glucose spillage (≥2+) in their urine at exam is compared with US Whites' rates of self-reported diabetes from the National Health and Nutrition Examination Survey, 1985.

among Hawaiian males (1.01 mmol/L [95% CI = 0.61, 1.55]) and especially in females (1.08 mmol/L [95% CI = 0.58, 1.58]) than values found for Whites of similar ages in the Lipid Research Clinics studies (males 1.16 mmol/L and females 1.49 mmol/L), indicating the possibility of substantially higher cardiovascular risk for Hawaiians.[11,12]

As shown in Figure 2, there was significant glycosuria (≥2+) in the spot urine or a history of the individuals being told they had an elevated blood sugar in 10 percent of females (95% CI = 5, 15) and 12 percent of males (95% CI = 6, 18). Among those 20–29, 30–39, 40–49, and 50–59 years of age, glycosylated hemoglobin levels greater than .11 were found in 2 percent, 7 percent, 8 percent, and 17 percent of these age groups, respectively, indicating an increased probability of glucose intolerance or poor control of diabetes.

Although there was no standardized follow-up, most of the six previously undiagnosed individuals with glycosuria are known to have been subsequently diagnosed by local physicians as having diabetes. The majority of the known diabetics were under poor control as evidenced by spillage of glucose in the urine and/or very high glycosylated hemoglobin levels.

Discussion

NHANES II found that 27 percent of US adults ages 20–59 were overweight,[9] compared with 65 percent of native Hawaiians in this survey, which used the same cut points for body mass index. The native Hawaiian males had more than twice the prevalence of overweight of US males in all age groups, whereas native Hawaiian females had 50 percent higher prevalence than US females. The predominantly upper-body type of fat distribution found in Hawaiians has been reported to be independently associated with higher risk of subsequent cardiovascular disease than is lower-body fat predominance.[10] However, ethnic differences in physiology and in the concept of ideal weight may contribute to the differences seen. Data from the 1985 NHIS indicate that 35 percent of US males and 31 percent of females in the age range 20 to 59 smoke.[7] The smoking rates among native Hawaiians appeared to be somewhat higher, and the age-specific patterns were most similar to US Blacks.

High rates of glucose intolerance have been found in some American Indian tribes.[13] Among urban Polynesians other than Hawaiians, rates of more than 12 percent for males and 11 percent for females have been reported, whereas rural rates are lower (1 to 4 percent).[14,15] In other Pacific populations, Micronesians from affluent, relatively westernized Nauru have a 30 percent prevalence of diabetes,[16] and Melanesian rates are low with rates in Fiji of 1.1 percent in the rural environment and 3.5 percent in an urban environment.[17] Many studies have implicated urbanization or westernization as a major risk factor for diabetes.[18]

The population studied was small, and previous experiences have made the Hawaiians suspicious of medical researchers; consequently overall response rates were held to 70 percent. For reasons of this kind, there are few population-based data on Hawaiians. In this study we had no indication that we were more or less successful in recruiting individuals with disease or elevated risk factors than healthier individuals, although the older individuals—who are more likely to have diagnosed disease and known risk factor elevations—were more likely to participate. As in any socioeconomically depressed population, out-migration is a problem and recruitment was somewhat more difficult in the young because of such factors as peer pressure, disillusionment, and work restrictions. As in any study, selective nonparticipation of healthy or ill individuals may have been possible. In order to better characterize the native Hawaiian population it is imperative that larger, better funded studies be carried out by working closely with native Hawaiians in multiple communities.

From these data, it appears that Hawaiians may be at high risk of premature coronary heart disease and stroke due to obesity, hypertension, hypercholesterolemia, low levels of HDL, smoking, and diabetes mellitus. In this population, individuals who were found to have significant abnormalities in the cardiovascular disease risk factors studied were often unaware of them, not under treatment, or not under adequate control. Such findings probably indicate serious deficiencies in health care systems and the knowledge of the population about the associated disease processes. ☐

Acknowledgments

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HIV Risk Behavior Reduction following Intervention with Key Opinion Leaders of Population: An Experimental Analysis

Jeffrey A. Kelly, PhD, Janet S. St. Lawrence, PhD, Yolanda E. Diaz, MSW, L. Yvonne Stevenson, MS, Allan C. Hauth, MA, Ted L. Brasfield, RTT, Seth C. Kalichman, BA, Joseph E. Smith, RN, and Michael E. Andrew, PhD

ABSTRACT

Background and Purpose. Peer norms influence the adoption of behavior changes to reduce risk for HIV (human immunodeficiency virus) infection. By experimentally intervening at a community level to modify risk behavior norms, it may be possible to promote generalized reductions in HIV risk practices within a population.

Methods. We trained persons reliably identified as popular opinion leaders among gay men in a small city to serve as behavior change endorsers to their peers. The opinion leaders acquired social skills for making these endorsements and complied in talking frequently with friends and acquaintances. Before and after intervention, we conducted surveys of men patronizing gay clubs in the intervention city and in two matched comparison cities.

Results. In the intervention city, the proportion of men who engaged in any unprotected anal intercourse in a two-month period decreased from 36.9 percent to 27.5 percent (−25 percent from baseline), with a reduction from 27.1 percent to 19.0 percent (−30 percent from baseline) for unprotected receptive anal intercourse. Relative to baseline levels, there was a 16 percent increase in condom use during anal intercourse and an 18 percent decrease in the proportion of men with more than one sexual partner. Little or no change was observed among men in the comparison cities over the same period of time.

Conclusions. Interventions that employ peer leaders to endorse change may produce or accelerate population behavior changes to lessen risk for HIV infection. (*Am J Public Health* 1991;81:168–171)

Introduction

Experience in many health promotion areas supports the importance of social influence as a determinant of population risk behavior patterns.[1-3] Gay men who successfully implement risk reduction report greater peer support for behavior change than their unsuccessful counterparts,[4-6] while beliefs that one's friends have already made precautionary changes and that these changes will be well accepted predict compliance with acquired immunodeficiency syndrome (AIDS) risk reduction recommendations.[5,6] Reductions in high-risk sexual behavior among homosexual men in urban centers appear linked to new norms in the gay communities of large cities which discourage high-risk activities such as unprotected anal intercourse and encourage precautionary steps such as condom use or other safer sex modifications.[7,8]

Persons frequently seek the advice of friends concerning steps needed to reduce risk for AIDS.[9] Diffusion of innovation theory posits that trends and innovations are often initiated by a relatively small segment of opinion leaders in the population.[10] Once visibly modeled and accepted by natural opinion leaders, innovations then diffuse throughout a population, influencing others. The purpose of this investigation was to experimentally test an AIDS prevention intervention based on diffusion of innovation/social influence principles. Because HIV (human immunodeficiency virus) risk behavior levels remain high among gay men in small cities[5,6,11] and because geographically isolated communities constitute compact environments where it is feasible to experimentally evaluate community-level intervention of this kind, the project targeted homosexual men in small cities.

Methods

Setting and Baseline Survey

The study was conducted late in 1989 in Biloxi and Hattiesburg, Mississippi, and Monroe, Louisiana, selected because they are relatively small (populations of 50,000 to 75,000 residents), separated by at least 60 miles from one another or from any other city of larger size, and because each city had one or two large gay bars. Due to the geographical isolation of the cities and the paucity of other meeting places, these clubs tend to attract large, stable crowds of homosexual men and serve as the primary social setting in each city's gay community. An earlier survey of men patronizing clubs in these cities revealed alarmingly high rates of high-risk sexual behavior.[6] AIDS incidence data for the cities are not compiled, but data from local health department HIV testing sites indicate relatively low seroprevalence rates (approximately 2 percent) among all men who seek testing in the cities. This suggests that HIV infection prevalence is still moderate but will increase unless behavior changes among gay men are quickly implemented.

From the University of Mississippi Medical Center, Jackson, MS. Address reprint requests to Jeffrey A. Kelly, PhD, Professor, Department of Psychiatry and Mental Health Sciences, Medical College of Wisconsin, 8701 Watertown Plank Road, Milwaukee, WI 53226. This paper, submitted to the Journal April 17, 1990, was revised and accepted for publication September 11, 1990.

To obtain baseline data on population AIDS risk characteristics, men entering each city's clubs were initially surveyed over three consecutive nights. The survey measure was completed anonymously and individually at small tables located near the club entrances; men who completed the questionnaire on one night did not complete it again during the same three-night sampling period. In addition to demographic information, the measure assessed:

- knowledge about AIDS risk behaviors and risk reduction steps using a 23-item true/false test[12];
- perceived social norms concerning the acceptability of adopting safer sex precautions measured by endorsement on 5-point Likert scales of five statements (example: "My gay male friends insist on using condoms when they have intercourse") and yielding scores of 5 to 25[5,6]; and
- personal sexual behavior occurring over the past two months, including number of occurrences and different partners for unprotected insertive and receptive anal intercourse, anal intercourse with condoms, and other low-risk practices (frottage, mutual masturbation, and similar activities). This measure was pilot tested extensively and has been successfully used in prior surveys of gay men's risk behavior.[5,6,11]

The same survey procedure was repeated in a second three-day sampling period four months later at each club in each city to establish baseline stability in population risk characteristics. A preliminary review of survey data revealed that AIDS risk knowledge scores were consistently high among men in each city ($M = 21$ of 23 items answered correctly) so basic knowledge items were deleted from subsequent surveys. Counts were made of the total number of men who entered clubs to determine the proportion of the entire population successfully surveyed.

Study Design and Intervention Procedures

Following baseline population assessments in all cities, one community (Biloxi) was randomly selected to receive the intervention while the other two communities served as comparison cities. The comparison cities received no specific intervention, although AIDS prevention posters and brochures were generally available in gay clubs in the cities. No significant differences in population risk behavior were found between the two comparison cities.

In the intervention city, a three-stage process identified, trained, and then contracted with popular opinion leaders among gay men to endorse behavior change to their peers.

Identification of Opinion Leaders: To identify key popular people, four club bartenders familiar with population members were trained to observe social interaction patterns within the city's gay clubs. Each made unobtrusive behavioral observations in the clubs for one week, recording first names and a physical identifier for 30 persons observed to socialize and be greeted positively most often by men in the clubs. Recording sheets were cross-matched for repeated mention; 36 of a total of 82 names received multiple nominations and were considered key popular people. Twenty-two were located, entered in the program, and asked to invite one friend also considered to be highly popular with gay men and not already a participant. The total training group consisted of 39 men and four women with a mean age of 30 years and a mean education level of 13.8 years; 91 percent of the group was White and 9 percent African-American or Hispanic.

Training the Opinion Leaders: The training consisted of four weekly, 90-minute group sessions led by a male and a female leader. In session one, the leaders reviewed basic epidemiology of HIV infection, high-risk behavior, and precautionary changes needed to reduce risk and misconceptions concerning risk. They discussed the steps that can be used to successfully implement risk reduction, including: keeping condoms readily available if sexually active; avoiding excessive intoxicant use before sex; discussing precautions in advance with sexual partners; resisting coercions to engage in high-risk practices; and self-reinforcing for behavior change efforts.

Session two described characteristics of effective health promotion messages: sensitizing others to the potential threat of AIDS; stressing that HIV infection can be prevented by behavior change; identifying specific behavior changes needed to reduce risk; suggesting strategies for implementing these changes; using self as an example in order to avoid a "preachy" tone (i.e. beginning statements with "I am learning to . . ." rather than "You should . . ."); and personally endorsing the positive value and desirable benefits of behavior change.

In the third session, leaders modeled conversational examples which incorporated the characteristics discussed in session two. Following discussion of the enactments, participants role-played how they would initiate conversations, incorporating these message characteristics. Group leaders and assistants observed this social skill rehearsal, providing feedback and suggestions. After becoming proficient in the role-plays, each participant identified four gay male friends with whom s/he could initiate an endorsement conversation in the next week. Monitoring forms were provided for recording descriptions of these conversations.

Session four reviewed outcomes of the real-life conversations; most participants reported positive reactions from others. The leaders and participants then problem-solved persons, settings, strategies, and times which would afford opportunities to initiate additional conversational contacts with other peers. All members agreed to initiate and monitor at least 10 more peer conversations over the next two weeks. The session concluded with further discussion emphasizing the participants' important role in stressing the benefits of behavior changes to prevent HIV infection among gay men in their own community. Of the 43 opinion leaders who began the intervention, 35 attended all the sessions.

Later review of participants' self-monitoring forms indicated that 371 peer conversations were monitored over the two-week period. This probably underestimates the number of conversations which actually took place because peer conversations were likely to have continued beyond the monitoring phase. In addition, before and after the training program, each opinion leader was asked to role-play standard, simulated peer educational conversations. Ratings of the tape recorded role-plays were made by trained judges, naive to whether practice conversations were pre- or postintervention. These rates revealed that, after training, the opinion leaders more frequently exhibited skills taught in the program (data available on request to author). These checks confirm social skill acquisition and conversational assignment compliance by the opinion leaders.

To stimulate opportunities for initiating peer conversations, posters with a traffic light logo (red, yellow, and green circles) but no printed explanation were placed throughout the clubs shortly before the end of the intervention. Each training participant was given a small lapel button with the same logo and asked to wear it. Participants reported that this cuing device led many people to inquire about the

TABLE 1—Measures of Behavior in Surveyed Populations Prior to Intervention and Behavior Change Postintervention

Behavior Characteristics[a]	Preintervention		3–6 Month Changes		Difference (95% CI)
	Intervention City	Comparison Cities	Intervention City	Comparison Cities	
Percent of Respondents Engaging in:					
Any Unprotected Anal Intercourse	36.9	40.4	−9.4	−2.9	6.5 (6.4, 6.6)
Unprotected Receptive Anal	27.1	30.9	−8.1	−3.7	4.4 (4.3, 4.5)
Unprotected Insertive Anal	31.5	31.5	−9.6	+0.4	10.0 (9.9, 10.1)
Percent of Anal Intercourse Occasions Protected by Condoms:					
Any Anal Intercourse	48.8	54.5	+6.5	−6.1	12.6 (12.5, 12.7)
Receptive Anal Intercourse	50.4	55.4	+8.0	−7.8	15.8 (15.7, 15.9)
Insertive Anal Intercourse	49.2	54.7	+5.2	−5.1	10.3 (10.2, 10.4)
Percent of Respondents Reporting Multiple Sexual Partners	42.4	38.2	−7.8	+1.9	9.7 (9.6, 9.8)

[a] In two months prior to survey point.

TABLE 2—Percentage Distribution for Number of Sexual Partners for Unprotected Anal Intercourse in Prior Two Months: Intervention and Comparison Populations Before and After Intervention

Number of Partners	Intervention City		Comparison Cities	
	Pre	Post	Pre	Post
0	70.7	80.0	68.8	70.9
1	7.1	5.2	7.9	8.4
2	4.2	3.4	3.5	3.7
3	2.0	1.0	4.0	3.4
4	1.9	1.3	6.5	6.6
5	2.6	0.9	2.5	4.0
6	2.9	1.2	1.3	0.8
7+	8.6	7.1	5.5	2.2

meaning of the symbol. These questions, in turn, stimulated opportunities to explain precautionary changes from the traffic light scheme (high-risk, moderate-risk, low-risk) and endorse these changes in the conversational style that was practiced in training. The lapel buttons also served as visual cues which reinforced the endorsement of behavior change by well-known and popular individuals.

Several months after the intervention's completion, participants were recontacted by telephone, encouraged to continue their peer education conversations, and commended for their continued efforts.

Determining Change in Population Risk Characteristics

Surveys of all male club patrons in all cities were repeated in the same manner as at baseline three months and again six months after the end of the training period. Responses of training participants were excluded. Intervention city men reported they were approached to initiate a mean of 7.1 conversations, demonstrating that the conversations encouraged in the training program occurred frequently.

Results

A total of 659 surveys were completed at the two baseline survey periods (intervention city N = 328, comparison cities N = 331) and 608 men completed surveys at the two postintervention periods (intervention city N = 278, comparison cities N = 330). This was an average of 81 percent (range = 68 to 88 percent) of all men patronizing each city's clubs during the survey periods. Between 24 and 47 percent of the men who completed measures on one occasion were the same men who completed them earlier; over 70 percent of respondents lived in the city where the survey was conducted. The mean age of all men surveyed was 29.1 years; the mean education level 14.9 years; 86 percent of men were White and 14 percent African-American or Hispanic. There were no significant differences between intervention and comparison city populations, nor were there systematic changes over time within a city's surveyed population, for these characteristics.

Table 1 presents behavioral characteristics for both the intervention and comparison cities before and following intervention with 95% CIs of the differences in behavior change. For purposes of brevity, results of the two baseline surveys and the three- and six-month postintervention surveys are averaged as a single figure. Populations in the intervention and comparison cities were initially comparable in risk characteristics. At the postintervention surveys, the proportion of men who engaged in unprotected anal intercourse decreased by about 30 percent from their initial levels. Much less change was evident in the population of men in the comparison cities. Similarly, there was an increase in proportion of all anal intercourse occasions when condoms were used only in the intervention city. Finally, the intervention produced a decrease in the proportion of intervention city men who reported having multiple sexual partners in the preceding two months, an effect not found in the comparison cities.

Table 2 shows the percentage distribution for number of sexual partners in the intervention and comparison populations before and after intervention. These results show some shifts reflecting an increased proportion of intervention city men who had either one or no sexual partner from pre- to postintervention and slight general decreases in the proportion of men who had two or more partners (Z

$= 1.66$, $p < .05$ for nonparametric Wilcoxon signed rank test).

Survey responses of the male opinion leaders, examined separately, showed that in the two months before training, 39 percent of opinion leaders engaged in unprotected anal intercourse; in the postintervention survey, 24 percent engaged in this practice.

Distinctive changes in social norm measure scores were not associated with the intervention; only a general and modest pre- to posttest increase in scores was found (from $M = 15.6$ to 17.0). However, intervention city men who refrained from unprotected anal intercourse reported in posttest surveys receiving health message conversations from peers more ($M = 6.3$ conversations) than men who still engaged in unprotected receptive anal intercourse ($M = 3.5$) or in any unprotected anal intercourse ($M = 4.8$).

Discussion

To the best of our knowledge, this constitutes the first report of an experimentally controlled study evaluating a community-level field test of an intervention for HIV risk reduction. This study found that engaging key opinion leaders popular with gay men in small cities to serve as behavior change endorsers to their peers, produced reductions in the proportion of men in the population who engaged in high-risk activities and produced concomitant, population-wide increases in precaution-taking. The magnitude of these shifts compares favorably to those commonly found following mass, community-based population intervention for other health behavior problems such as cigarette smoking and cardiovascular risk reduction. This suggests a cost-effective HIV prevention model which may be of benefit in interventions with gay men, intravenous drug users, adolescents, and other groups.

Several findings in this study were of particular interest. First, AIDS risk behavior knowledge scores were high among men in each city even before intervention, suggesting that lack of knowledge about AIDS was not a critical dimension in the populations studied and that intervention methods which induce behavior change implementation and social acceptability of changes are needed. Second, while we anticipated that the intervention would alter perceptions of social norms, the norm measure did not reflect changes parallel to the population behavior shifts. Generalized changes in sexual behavior may precede peer norm changes rather than cause the initial adoption of risk reduction. In that case, elements of the intervention responsible for population behavior change may have included frequent safety prompts delivered by credible peers, conversational dissemination of suggestions for implementing precautionary changes, legitimization of making risk reduction changes by opinion leaders, or other factors.

While the study's results are highly promising, a number of cautions must be raised. We studied only men who patronized clubs and relied on self-reports of behavior. However, the geographical isolation of the study cities, the stability of club populations over time, and the function of clubs as the major social setting for gay men in these communities supports the relevance of data collection at these sites. While self-report behavior measures are essential in large-scale HIV risk behavior research, other data corroborating population behavior change—such as changes in community STD (sexually transmitted disease) incidence, condom purchasing patterns, or HIV testing requests and infection incidence—are needed to substantiate future conclusions about intervention impact. Finally, only a small number of cities were involved in this demonstration. Wider-scale clinical field trials of such experimental models in multiple cities and with larger populations are needed to establish their generalizability. ☐

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Motorcycle Licensure, Ownership, and Injury Crash Involvement

ABSTRACT

The interrelationships among motorcycle licensure, ownership, and injury crash involvement were investigated in a sample of 2,723 motorcycle drivers severely or fatally injured in California in 1985–86. Owners of motorcycles in such crashes ("driver-owners") were less likely to have valid licenses than a random sample of motorcycle owners who had not been in crashes (42 vs. 57 percent). Thirty-three percent of the crash-involved drivers had valid motorcycle driver's licenses; 39 percent were operating motorcycles they did not own ("driver-nonowners"). Driver-nonowners were less likely to be validly licensed than driver-owners (20 percent vs. 44 percent). The licensing rate of crash-involved driver-nonowners was 15 percent if the owner was also unlicensed. Rates of valid licensure were lowest among the youngest drivers. Virtually no crash-involved driver-nonowners under age 21 were licensed in cases in which the owner was also young and unlicensed. (*Am J Public Health* 1991;81:172–176)

Jess F. Kraus, PhD, Craig Anderson, DHSc, Paul Zador, PhD, Allan Williams, PhD, Shakeh Arzemanian, MPH, Weichang Li, MA, and Michael Salatka, MPH

Introduction

In most states a special license is required to operate a motorcycle on public streets and highways. There is evidence, however, that many of the motorcycle drivers killed or seriously injured in crashes did not have a license to operate a motorcycle. Hurt and associates* reported in an unpublished document that 46 percent of 900 motorcycle drivers involved in crashes in Los Angeles in 1976–77 were not validly licensed. Data from the US Department of Transportation's Fatal Accident Reporting System (FARS)** show that about 43 percent of motorcyclists killed on public roads or highways in the United States did not hold a valid motorcycle operator's license. In both reports, however, there was no information given on licensure status of motorcyclists not involved in injury crashes. Neither of these data sets provides information on whether the injured were driving motorcycles they owned, nor do they indicate whether lack of licensure was associated with increased crash risk. Barry[1] noted 20 years ago that a substantial percentage of operators of motorcycles involved in injury crashes had "borrowed" the vehicle, but the report did not include information on whether the motorcyclists were licensed.

The present study was designed to address the following questions:

- Do motorcycles owned by people without motorcycle licenses have greater likelihood of crash involvement?
- To what extent do motorcycle drivers in serious and fatal injury crashes have valid licenses? own the motorcycle they were driving?
- How does motorcycle licensure vary by whether the crash-involved driver owned the motorcycle ("driver-owner" vs. "driver-nonowner")?
- How is motorcycle licensure affected by the relationship between crash-involved driver-nonowners and owners, when other relevant factors are also controlled?

Methods

A valid motorcycle license in California is a "class 4" license or, more frequently, a motorcycle endorsement to another class of license. Motorcycle licensure requires a written examination and driving test for two-wheeled vehicles. Motorcyclist training was not required for any drivers prior to 1988.[2]

Crash-Involved Motorcycles

Motorcycle crashes that occurred between July 1, 1985 and October 31, 1986

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*Hurt *et al*: Motorcycle accident cause factors and identification of countermeasures. Traffic Safety Center, University of Southern California, Los Angeles, 1981.

**Buchanan L: Presentation at Transportation Research Board Meeting, Washington, DC, 1988.

Address reprint requests to Jess F. Kraus, PhD, Department of Epidemiology, School of Public Health, UCLA, Los Angeles, CA 90024-1772. Co-authors Anderson, Arzemanian, Li, and Salatka are also at UCLA. Dr Zador and Dr Williams are with the Insurance Institute for Highway Safety, 1005 N. Glebe Road, Arlington, VA 22201. This paper, submitted to the Journal February 28, 1990, was revised and accepted for publication June 14, 1990.

TABLE 1—Availability of Information on License Status and Percent Validly Licensed Motorcyclists among Owners and Drivers of Motorcycles in Study Group

Driver/Ownership Status	Total Number	Driver License Information Number (%) with Known License Status	% with Valid License for Operating a Motorcycle among Those with Known License Status
MOTORCYCLES IN CRASHES			
(a) Driver-owner	1,467	1,438 (98)	44
(b) Driver-nonowner	930	875 (94)	20
(c) Owner-nondriver	930	525 (56)	35
(d) Driver-unknown if owner	326	241 (74)	10
All drivers (a,b,d)	2,723	2,554 (94)	33
All owners (a,c)*	2,397	1,963 (82)	42
MOTORCYCLES NOT IN CRASHES			
Owners	13,638	8,648** (63)	57

*Excludes 326 cases in which ownership was undetermined.

**Subsequent results based on 6,965 motorcycle owners with license records (see Methods section).

on a California road, street, or highway were identified using the California Highway Patrol (CHP) Statewide Integrated Traffic Records Systems (SWITRS) and the original police reports. Crashes analyzed were limited to those in which the motorcycle driver was fatally or severely injured. A severe injury was defined by the CHP[3] as a cut or laceration of skin from which blood flows freely or saturates clothing, and also includes an extremity not in normal position, indicating fracture. In almost all cases, the injured driver was taken to a hospital by an ambulance or paramedic unit.*** Mopeds, scooters, all-terrain vehicles, other off-road vehicles, motorcycles of 1980 or earlier model years, and motorcycles with out-of-state registrations were excluded.

The SWITRS and the police reports provided data about the driver (driver's license number, age, sex, and extent of injury); the vehicle (vehicle license number, identification number—when recorded—make, and model year), and name and address of the registered owner.

Crash-Involved Drivers: The license status of injured drivers was determined from the California Department of Motor Vehicles (DMV) driver history file using the driver's license number recorded on the crash report. Drivers not licensed to operate motorcycles were classified as having a valid license for another class of vehicle, having a suspended or revoked license, or having no license at all.

Unlicensed operators included the following categories: no driver's license number recorded on the crash report, those whose license expired before the crash date, those whose license was first issued after the crash date, driver's license number beginning with "X" (indicating a DMV file with information on violations and crashes but no driver's license issued), and those with the DMV identification card only.

License status was classified as unknown if the driver was licensed in another state or country, if the driver was listed on the crash report as having an instruction permit or a temporary license, or if no computer record could be located for the driver's license number written on the crash report. If the driver's license history contained no record of a crash on the date of interest and the driver's birth date was inconsistent with the age on the crash report, the driver's license history data were not used and license status was also considered to be unknown.

Owners of Crash-Involved Motorcycles: In cases in which the crash-involved driver did not own the motorcycle, information on the license status of the owner was obtained. Because the owner's name and addresses—but not the owner's driver's license number—are on the crash report, it was necessary to match the name and address information on the crash report with name and address information on the California driver's license history file to identify the owners' driver's license number and obtain licensure information.

Owners of Motorcycles Not in Crashes (Comparison Group)

A comparison group composed of a random sample of 13,638 motorcycles, representing 4.5 percent of all registered motorcycles in California, was selected from the DMV vehicle registration file as of November 1, 1986. Inclusion in this sample was restricted in the same way as for motorcycles involved in crashes.

The chance of a motorcycle being included in the sample was proportional to the number of months in which it was validly registered in California during the study period. Once in the sample, motorcycles were assigned with equal probability to each of the 16 months in the study during which they were registered for use. This procedure resulted in a comparison sample that included every motorcycle that was registered during a month with equal probability for that month.

The driver's license history for the owners of motorcycles in the comparison sample was obtained in the following way. The status of the license was determined for the reference month to which the motorcycle was assigned for comparison. Of the initial sample of 13,638 motorcycles, name and address for the registered owner were available for 12,481 of them (92 percent). The driver's license number was located for 8,669 of the 12,481 (69 percent). Of the 8,669, a driver's license history file was located for 8,648 owners (99.8 percent). Based on the vehicle identification number, 8,510 of the vehicles were found to be true motorcycles. Driver's license histories of owners were randomly discarded until all months had a sampling ratio of 2.56 controls per case, the ratio found in the sparsest month. The final comparison sample included 6,965 vehicle owners.

Results

Licensure Information

The information available for subgroups of motorcycle owners and drivers is provided in Table 1. Licensure information was available for 94 percent of crash-involved drivers but for only 56 percent of owners of motorcycles in crashes who were not driving the motorcycle and for 63 percent of the comparison sample of owners of motorcycles not in crashes.

Licensure Rates

Table 1 also provides a summary of motorcycle licensure rates for the various groups of motorcycle owners and drivers.

***Personal communication, California Highway Patrol, April 1990.

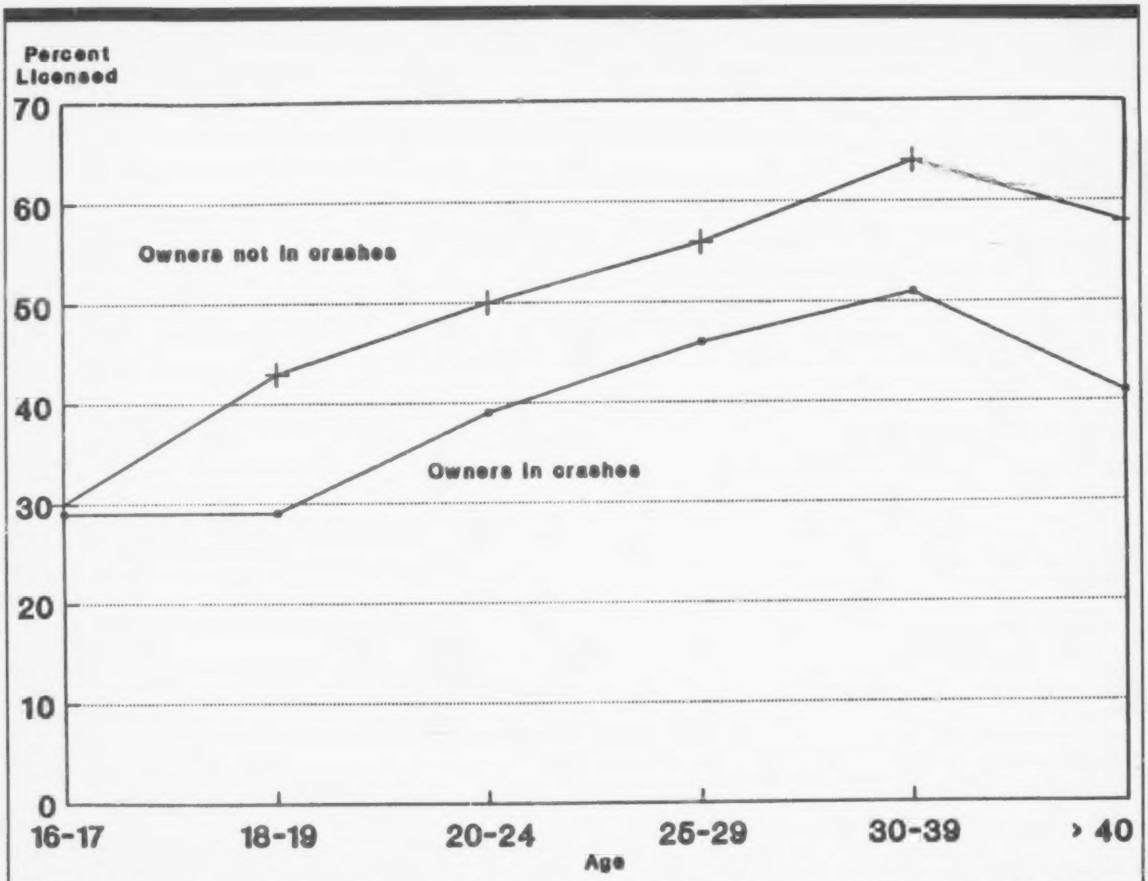

FIGURE 1—Percent of Validly Licensed Motorcyclists: Owners of Motorcycles in Crashes vs. Owners of Motorcycles not in Crashes, by Age

Owners: Owners of motorcycles in crashes were less likely to have valid licenses than motorcycle owners not in crashes (42 percent vs. 57 percent, odds ratio [OR] = 0.53, 95 percent confidence interval [CI] = 0.48, 0.59). Owners of crashed motorcycles were younger than the comparison group of motorcycle owners: 44 percent were under age 25 compared with 22 percent of the comparison group (OR = 2.86, CI = 2.58, 3.18). Owner licensure rates in both groups increased with age until age 40; fewer owners age 40 and over were licensed than owners ages 30-39. Figure 1 shows that the younger age of owners of crashed motorcycles did not account for the relationship between owner licensure and crash involvement. At all ages, owners of crashed motorcycles were less likely to be validly licensed than owners not in crashes (OR = 0.60, CI = 0.54, 0.66).

Owners of motorcycles in crashes who were operating the motorcycle were also less likely to be validly licensed than motorcycle owners not in crashes (44 vs. 57 percent, OR = 0.59, CI = 0.53, 0.66).

Drivers: Overall, 33 percent of crash-involved drivers were validly licensed. Valid licensure was strongly associated with age. The lowest licensure rate was among those under age 20 (18 percent); the highest rate was among those age 40 and older (51 percent).

Thirty-nine percent of the drivers were operating motorcycles they did not own, and they were less likely to be licensed than driver-owners (20 percent vs. 44 percent, OR = 0.31, CI = 0.26, 0.38). Driver-nonowners tended to be younger than driver-owners (63 percent were under age 25 compared with 47 percent of driver-owners; OR = 1.97, CI = 1.65, 2.34), but as shown in Figure 2, the differences in licensure rates between these two groups held, regardless of age (OR = 0.35, CI = 0.29, 0.43).

Among crash-involved drivers without valid licenses, 53 percent had a license to drive a passenger car but did not have a motorcycle endorsement, 31 percent did not have a license to drive any kind of motor vehicle, and 16 percent had suspended or revoked licenses.

Owner-Driver Pairs

Table 2 displays the percentage of validly licensed driver-nonowners by age, motorcycle owner license status, and the relationship between the driver and the owner. Older owners are defined as those who are more than 10 years older than the driver. Overall, in 56 percent of the 517 pairs, neither driver-nonowner nor owner was validly licensed. Valid driver licensure increased with driver age and with having a licensed owner. When owners were licensed, 27 percent of the drivers held valid licenses. When owners were unlicensed, only 14 percent of the drivers were licensed.

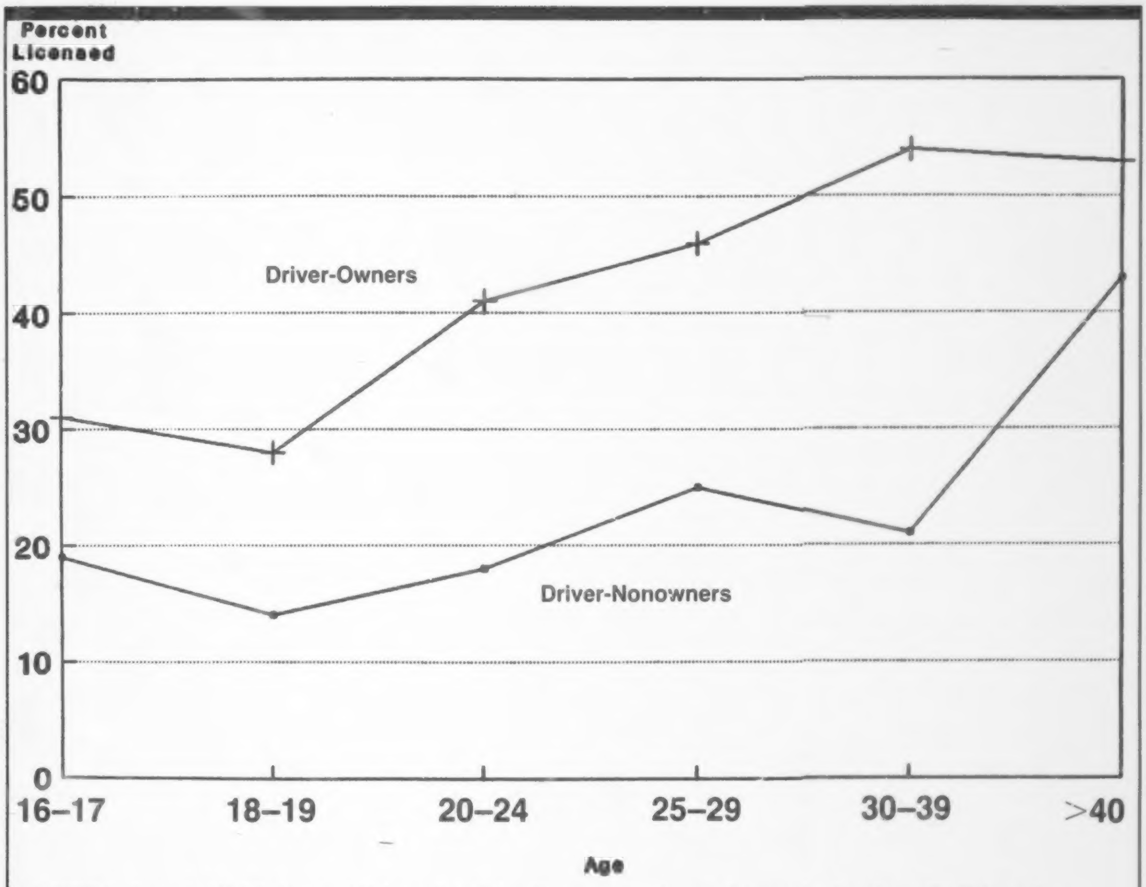

FIGURE 2—Percent of Validly Licensed Motorcyclists: Driver/Owners in Crashes vs. Drivers who do not own the Motorcycle

The data were analyzed using a log-linear model that included the three factors shown in Table 2 plus the interaction between motorcycle owner license status and the relationship between the driver-nonowner and the motorcycle owner. A greater proportion of drivers were licensed to operate a motorcycle if they were older and if the owner was licensed. More drivers were licensed if the owner was at least 10 years older than the driver and had the same last name or address as the driver, and fewer drivers were licensed when the owner was within 10 years of the driver's age and had a different last name and address.

Discussion

The proportion of unlicensed injured motorcycle drivers (67 percent) identified in this study is extraordinarily high compared with the proportion of fatally injured motor vehicle drivers (11 percent) in the US Fatal Accident Reporting System (FARS).[4] Our results also indicate that motorcycles owned by people without motorcycle licenses are more likely to be in crashes than motorcycles owned by people with valid motorcycle driving licenses. The results strengthen earlier, incomplete findings by Hurt and associates and FARS that most motorcyclists in crashes, especially young motorcyclists, are not validly licensed. In addition, those driving motorcycles they did not own were likely to be unlicensed, especially if the owner was also not licensed to operate a motorcycle.

The results also show that driver age, the use of someone else's motorcycle, and borrowing from an unlicensed owner are independent factors affecting valid licensure. Among young motorcyclists who crashed on motorcycles that were owned by other young and unlicensed motorcyclists, virtually none was validly licensed.

Unavailable data may alter somewhat the percentages presented. Licensure data were available for most of the drivers, but the driver's license status of only about 60 percent of the owners of motorcycles crashed by another operator and the random sample of owners of motorcycles not in crashes was known. This occurred usually because the name and address of the owner could not be matched with a name and address on the driver's license file. This linkage can fail for various reasons, such as a change of address by the owner. It is not known whether owners who could not be traced had a different percentage of valid licensure than those successfully traced.

Several countermeasures can be considered for reducing the crashes associated with invalidly licensed motorcycle owners and drivers. Requiring proof of valid licensure as a prerequisite for purchasing a motorcycle would reduce the extent to which owners and drivers are not

TABLE 2—Percent of Motorcycle Driver-Nonowners Validly Licensed by Owner Relationship to Driver, Driver Age, and Owner License Status

Owner Relationship to Driver	Owner Licensure Status	Driver's Age		
		16–19	20–24	25+
Age within 10 years	Licensed	33[a]	35	50
Same name/address	Unlicensed	12	13	13
Age within 10 years	Licensed	5	18	13
Different name/address	Unlicensed	0	5	14
>10 years older than driver	Licensed	33[a]	67[a]	67[a]
Same name/address	Unlicensed	28	53	33[a]
All others[b]	Licensed	67[a]	11[a]	12
	Unlicensed	6	6	25

[a] Based on fewer than 10 observations.
[b] All other driver-nonowner and owner pairs.

properly licensed. Motorcycle driving by invalidly licensed people might be reduced by stepping up enforcement of licensure laws and/or providing more stringent penalties for lack of licensure. Penalties against the owners of motorcycles driven by invalidly licensed drivers could deter owners from allowing such people to operate their motorcycles.

If applied, these measures could change the behavior of motorcycle owners and operators in one of two ways: these people can avoid sanctions either by obtaining the proper license or by not driving.

Measures that merely force licensure but do not change the amount of driving may have no effect on crashes. When existing licensure and training programs have been formally evaluated based on random assignment, the programs have not been shown to be effective in reducing crashes.[†,††] Thus, sanctions applied to unlicensed drivers and owners are more likely to be effective if they reduce driving than if they merely shift unlicensed driving to licensed driving. If the measures reduce exposure, they will reduce crashes, especially of young, novice, unlicensed drivers on borrowed bikes. □

Acknowledgments

This work would not have been possible without the splendid cooperation of the Department of the California Highway Patrol Commissioner, James E. Smith; Management Information Section (Sgt. John Bailey and Bev Christ); California Department of Motor Vehicles, Division of EDP Service (Neal Graham, Associate Data Processing Analyst); Drivers License Automation Development Unit (Rudy Modelo); and Research and Development Section (William Marsh). We also thank Steven Harrington and Ching-Sung Wu for computer-programming assistance.

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Alcohol Use and Abuse in Random Samples of Physicians and Medical Students

William E. McAuliffe, PhD, Mary Rohman, PhD, Paul Breer, PhD, Grace Wyshak, PhD, Susan Santangelo, BA, and Elizabeth Magnuson, MS

ABSTRACT

Background. This study sought to resolve conflicting views about whether physicians are especially prone to alcohol abuse.

Methods. Using an anonymous, mailed questionnaire on substance use, we surveyed 500 physicians, 510 pharmacists, and 974 of their students. The physicians and pharmacists were selected randomly from the state society's membership lists, and students selected were from local school lists. Follow-up surveys were sent to nonresponders at two-week intervals.

Results. The physicians and medical students did not drink especially heavily and were no more vulnerable to alcoholism than were their counterparts in pharmacy and other professions. Physicians differed from pharmacists in their style of drinking (greater frequency, smaller quantity), but not in total amount of alcohol consumed. Drinking habits among physicians were not associated with medical specialty or type of practice, but were positively related to gender (males drank more than females) and to age (older doctors were more apt to qualify as heavy drinkers than were younger doctors).

Conclusions. Physicians were no more likely to abuse substances nonmedically than were other professionals. Any group in which alcohol use is nearly universal incurs a risk of abuse and impairment that cannot be ignored. (*Am J Public Health* 1991; 81:177–182)

Introduction

The authors previously presented evidence that many physicians, pharmacists, and their students have used controlled drugs for self-treatment and nonmedical purposes.[1] In his commentary, Lewis[2] cited the authors' failure to include alcohol use in their report. Some authors[3–5] have concluded that physicians are especially prone to alcoholism, while others[6–9] have concluded the opposite. Brewster[6] identified methodological flaws and contradictory findings in studies comparing physicians to other groups regarding cirrhosis mortality[10–13] and hospital admissions for alcoholism.[14–19] She concluded that surveys of physicians regarding alcohol use produce the most valid comparisons.[6,*] Our search of the literature revealed that few surveys have been reported to date,[7,8,20–25] and these have been limited by either small,[7,20,21] selective,[20,23] or nonrandom samples;[8,21,24] lack of anonymity for respondents;[21] incomplete measures;[27] minimal reporting of alcohol results;[22] or absence of comparable data on nonphysicians.[7,24,25] Studies of medical students suffered from similar limitations,[26] but recent studies have improved our knowledge substantially.[7,27–34]

This article presents a comprehensive picture of current and lifetime physician alcohol use and abuse from surveys of doctors and medical students, with comparative data on pharmacists, pharmacy students, and lawyers. The article also describes how physicians' drinking varies by medical specialty, type of practice, gender, and age.

Methods

The bulk of the data come from a previously described,[1] mailed, questionnaire survey (November 1984) of 500 physicians randomly sampled from the membership (8,400) of a state medical society, and 504 medical students randomly drawn from medical schools (combined enrollment of 1,438) in the same state. Approximately 73 percent of the state's physicians were society members in 1984. Comparative data come from a sample of 510 pharmacists randomly selected from the membership (1,270) of the state's pharmaceutical society and 470 pharmacy students randomly drawn from area schools (combined enrollment of 1,160). Each mailing packet contained a questionnaire, a letter assuring anonymity, and a post card with instructions to return it separately so that names could be removed from the follow-up mailing list. Second and third mailings and a reminder phone call to nonrespondents followed at two-

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[*] McAuliffe WE, Breer PE: Are physicians more susceptible to alcoholism and substance abuse than are other professionals? The con view. Paper presented at the Joint Meeting of the American Medical Society on Alcoholism and Other Drug Dependencies and the Research Society on Alcoholism, San Francisco, CA, April 22, 1986.

Address reprint requests to William E. McAuliffe, Department of Psychiatry, Harvard Medical School, Cambridge Hospital, 1493 Cambridge Street, Cambridge, MA 02139. He is also with the Department of Behavioral Sciences at Harvard School of Public Health; Dr. Breer also holds an appointment at the Harvard Medical School. Dr. Rohman is with Blue Cross Blue Shield of Massachusetts; Dr. Wyshak is with the Center for Population Studies, Harvard University; Ms. Santangelo is with the Department of Epidemiology, and Ms. Magnuson is with the Department of Biostatistics, both at Harvard School of Public Health. This paper, submitted to the Journal June 22, 1990, was revised and accepted for publication September 20, 1990.

TABLE 1—Drinking in the Last Year

	Physicians (n = 337)	Pharmacists (n = 312)	Medical Students (n = 381)	Pharmacy Students (n = 278)
Quantity/Frequency				
Mean drinks/month[a]	20.3	21.2	13.3[#]	15.1[*]
Mean drinking days/month	9.9	8.1[*]	5.3[#]	3.9[#]
Mean drinks/drinking day	1.7	1.9[*]	2.1[#]	2.9[*#]
Admitted Having Current Drinking Problem (%)	2	1	3	1

[a] Exceeds product of drinking days per month and drinks per day. The mean of ratios (products) equals the ratio of means only under special conditions.[38]

[*] Difference ($p < .05$) either between physicians and pharmacists or between medical students and pharmacy students, controlling for age, gender, and citizenship.

[#] Difference ($p < .05$) either between physicians and medical students or between pharmacists and pharmacy students, controlling for gender and citizenship.

week intervals. The Committee on the Use of Human Subjects in Research at the Harvard School of Public Health approved the procedures.

Measures of Drinking Behavior and Problems

Responding to fixed alternative items, subjects qualitatively characterized their drinking since entering college, recorded whether they currently or ever had a drinking problem, and how frequently they experienced specific dysfunctions due to drinking (Table 2). For analysis we divided dysfunctions into "minor" and "major." Subjects also reported how often they drank in the past year (frequency) and how many drinks they typically consumed on each occasion (quantity). One drink was defined as 12 oz of beer, 3.5 oz of wine, or 1 oz of liquor. From those quantity/frequency reports, the investigators constructed a typology of current drinking patterns using the cutoff for heavy drinking (61+ drinks per month) employed by other researchers.[35,36]

Alcohol Abuser and Potential Abuser Operationally Defined

"Alcohol abuse" included both pathological use and significant dysfunction due to drinking, in accordance with DSM-III criteria.[37] "Pathological use" meant an average of more than three drinks per day in the past year, characterizing self as a lifetime "problem" drinker, or admitting a past or current drinking problem. "Significant dysfunction" was operationalized as a score of at least three on a scale of *minor* dysfunctions *or* a score of at least one on a scale of *major* dysfunctions. Scores were computed by assigning zero for a response of "never," one for "once or twice," two for "sometimes," and three for "often," and then summing across the four items in each scale.

A "potential abuser" had to report a combination of "heavy" alcohol use and "some dysfunction" due to drinking. "Heavy use" meant either an average of more than two drinks per day in the past year, or self-characterization as a lifetime heavy drinker. "Some dysfunction" was defined as at least one minor or major dysfunction due to drinking.

Statistical Methods

The methods used for comparing sample statistics included 95 percent confidence intervals and significance tests based on the t and chi-square distributions, multiple regression, and multiple logistic regression, depending on the number and measurement level of variables. The standard error calculations assumed infinite populations; all tests were two-tailed.

The Wyshak Sample

Unpublished quantity/frequency data will also be presented from a second survey of physicians and lawyers described elsewhere by Wyshak, *et al.*[22] This second physician sample was drawn randomly from the Department of Public Health's list of registered physicians in the same state five years before.

Results

Survey Response

After unlocatable and deceased respondents were removed from the four samples, the response rates were 70 percent for physicians, 78 percent for medical students, 76 percent for pharmacists, and 67 percent for pharmacy students. The figures for the two professional samples compared favorably with Wyshak's survey.[22] A few questionnaires returned by nonpracticing professionals and students who had left school were excluded to produce final samples of 337 physicians, 312 pharmacists, 381 medical students, and 278 pharmacy students. Responding physicians did not differ significantly from nonrespondents with respect to specialty, type of practice, sex, or geographical distribution. Only data on sex were available for the other three samples; medical school females responded more often than males (83 percent vs 75 percent, respectively), but no difference was found among pharmacists or pharmacy students. For the survey as a whole, how soon a respondent returned his/her questionnaire did not correlate with number of drinks in the past year, lifetime pattern of drinking, or reporting having ever had a drinking problem. However, doctors who reported having ever had a drinking problem tended to return their questionnaires later rather than sooner ($r = .09$; CI = $+.04, +.14$).

Despite the educational differences between the physicians and pharmacists, the samples were nearly identical in age (mean of 46 for both), gender (85 and 84 percent male respectively), and US citizenship (97 and 99 percent respectively). More pharmacists (79 percent) than physicians (66 percent) attended religious services. Medical students were older (25 vs 22 years), more often males (59 vs 44 percent), less religious, and more often American (99 vs 91 percent) than were pharmacy students. When appropriate, these variables are controlled in subsequent analyses, but since the samples were so similar, the adjustments made a difference only between practitioners and students (age was not controlled).

Drinking in the Last Year

There is no consistent evidence that physicians or medical students currently drink more than their counterparts in pharmacy (Table 1). While physicians and medical students reported drinking more *frequently* than the pharmacy samples, the pharmacy samples reported consuming greater *quantities* per drinking episode. In both medicine and pharmacy, practitioners drank more often than students but consumed less per episode. These zero-order differences remained when demographics were controlled.

There were only minor differences in current patterns of drinking among the four samples. Most respondents (e.g., 86 percent of doctors) in all four groups were

TABLE 2—Dysfunction Due to Drinking (% Ever)[a]

Has your drinking ever caused you to:	Physicians (n = 337)	Pharmacists (n = 312)	Medical Students (n = 381)	Pharmacy Students (n = 278)
Minor Dysfunction				
Get behind in your work/studies	1	4	9[c]	12[c]
Call in sick or be late for work/school	4	5	15[c]	14[c]
Have trouble getting along with people	4	3	6	4
Worry that you might be using too much or too often	15	8[b]	13	9
One or more minor dysfunctions	17	11	22	21[c]
Major Dysfunction				
Seriously consider suicide	1	0	1	1
Have an auto accident or other type of accident	2	3	4	2
Provide less than your best patient care/get poor grades	3	2	2	6
See a psychiatrist, psychologist, social worker, or counselor	3	2	1	1
One or more major dysfunctions	5	5	5	8

[a] Item response options were "never," "once or twice," "sometimes," and "often." All percentages here are for "once or twice" or greater.

[b] Difference (p < .05) between either physicians and pharmacists or medical students and pharmacy students.

[c] Difference (p < .05) between either physicians and medical students or pharmacists and pharmacy students.

either light or moderate drinkers. The primary difference between the medical and pharmacy samples was in the percentage of abstainers (4 percent of doctors, 5 percent of medical students, 10 percent of pharmacists, and 11 percent of pharmacy students). Currently abstaining practitioners had a past or have a current (nonactive) drinking problem in some cases (one-third of physician abstainers, one-seventh of pharmacist abstainers). The major difference between practitioners and students was in the percentage of heavy drinkers (10 percent doctors, 12 percent pharmacists, 6 percent both medical and pharmacy students). When the amount of drinking was dichotomized at 61 drinks per month (cutoff for heavy drinking), the percentage difference between practitioners and students was estimated by logistic regression to be 5.7 (95 CI = 2.4, 8.9); that difference declined to 3.4 (0.1, 6.8) when gender and citizenship were included.[39]

Few respondents admitted to having a current drinking problem (from 1 to 3 percent; Table 1). There was, however, evidence of possible denial among heavy drinkers. While most of the respondents who admitted to a drinking problem also reported drinking very heavily (91+ drinks per month), the reverse was not true. Of the 13 physicians who reported drinking very heavily, only four saw themselves as having a current drinking problem, even though some of them admitted averaging over six drinks a day. Among pharmacists, only two of 17 very heavy drinkers reported a current problem. Among students, the figures were five of nine in medicine and zero of seven in pharmacy.

The Wyshak, *et al*, sample reported similar levels of drinking in the last year. With females eliminated to facilitate comparison, the mean of drinks per episode was almost identical (1.8 vs 1.7), and slightly more of Wyshak's sample reported heavy drinking (15.6 vs 12.4). Wyshak's physicians, however, averaged more drinking days per year (152 vs 126) than did ours (p < .05).

Lifetime Drinking Behavior and Problems

When asked to characterize their drinking pattern since entering college, most respondents (e.g., 91 percent of physicians) also described themselves as either light or moderate drinkers. Physicians differed significantly from pharmacists primarily in the percentage of abstainers (6 percent doctors, 12 percent pharmacists) as did medical students (7 percent) from pharmacy students (14 percent). There were no differences across the samples in the percentage of lifetime heavy or problem drinking. In all four samples, 4 percent of respondents admitted having ever had a drinking problem since entering college.

Dysfunction Due to Drinking

Although one in six physicians had worried at least once that he or she might be drinking too much, few reported a major dysfunction due to drinking (Table 2). Three percent of the physicians indicated that at least "once or twice" drinking had caused them to give less than their best patient care, and 3 percent had sought professional treatment for their drinking at some time in their careers. All told, 17 percent of physicians had ever experienced at least one dysfunction, while 10 percent had experienced two or more.

Physicians were more apt to worry about their drinking then were pharmacists (ratio of 1.9:1), and students were more likely than practitioners to report at least one minor dysfunction (ratio of 1.8:1). Most of the difference between practitioners and students, however, was in the two items that were worded differently for the two groups (work vs school).

Lifetime Abusers and Potential Abusers

Although physicians and medical students had slightly smaller percentages of combined abusers and potential abusers than the pharmacy samples (Figure 1), none of the samples differed significantly from the others on abuse, potential abuse, or the sum of the two. Controlling for gender and citizenship did not alter that finding. Of the 4 percent of physicians who abused alcohol at some time in their lives, half reported either no drinking or moderate drinking in the past year; the other half were still drinking heavily.

Subject Characteristics and Alcohol Use

Alcohol use was not related to medical specialty, type of practice (solo, group, and so on), citizenship, or religiosity, but was associated with both gender and age. Male physicians averaged 10.5 drinking days per month in the last year to only 6.5 for female physicians (CI for difference = 1.2, 7.0); male physicians also consumed more drinks per occasion (1.7 vs 1.2; CI for difference = 0.2, 0.8). However, like the male physicians, 4 percent of the female physicians qualified as ever being alcohol abusers, primarily on the basis of having a past drinking problem.

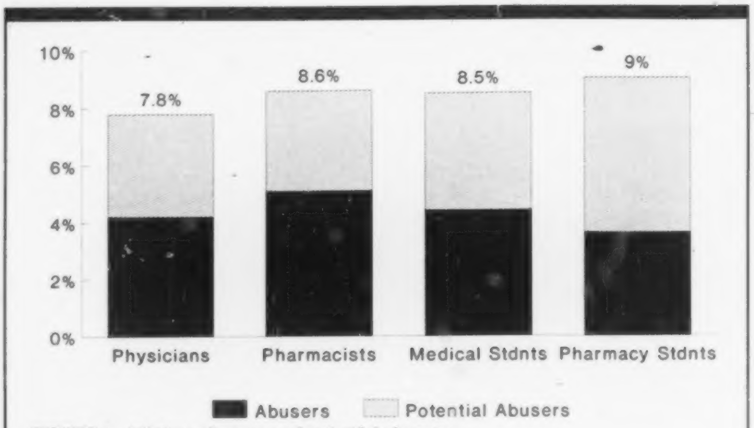

FIGURE 1—Lifetime abusers and potential abusers.

Abusers = pathological use plus impairment. Potential Abusers = pathological use plus minor impairment.

The gender difference in current drinking behavior was equally striking among pharmacists, but not among the student groups. Among medical students, males drank more than females in the last year, averaging 15.2 drinks per month to 10.4 for the females (CI for the difference = 0.8, 9.0). The sex difference, however, was only about a third as large as that for physicians.

Whereas gender bears the same relationship to drinking among our four samples as it does in the US population as a whole, the relationship between age and drinking is reversed (Figure 2, where females have been removed from all samples). Heavy drinking *decreases* with age for US males in general[35] (a finding confirmed in cross-sectional and cohort studies[40]), but it *increases* with age for both physician samples. Like physicians, pharmacists begin drinking much less than do US males, but jump quickly to a level close to that of the US population and remain at that level thereafter. Lawyers start somewhat lower than US males, peak sharply in their forties, and decline to a level slightly higher than physicians in their fifties and sixties.

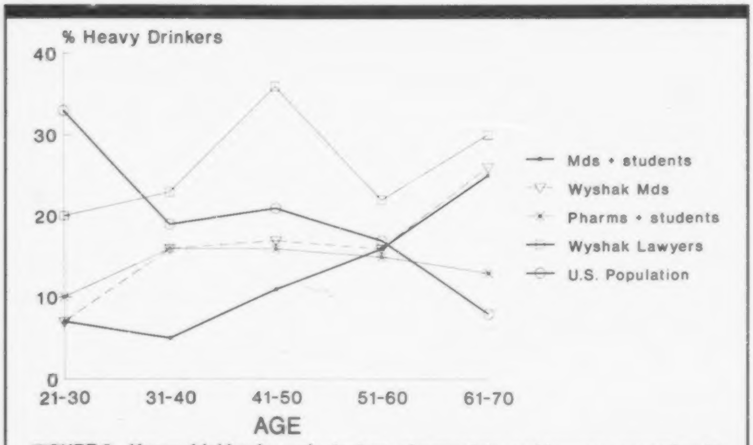

FIGURE 2—Heavy drinking by male doctors, pharmacists, and lawyers, according to age.

Note: US data from Clark and Midanik.[35]

Discussion

None of the present findings from our sample or the Wyshak, *et al*, sample indicates that physicians or medical students are especially prone to heavy drinking, alcohol problems, or alcohol abuse. The unique relationship between age and heavy drinking in these medical samples is confirmed by small surveys of residents and young physicians by the present authors[7] and another team in the same state,** and by physicians in another country.[25] It was also supported by data from a 20-year longitudinal study comparing a small sample of physicians with controls.[20] Analysis of data on job and life stress from this present study and our previous investigation[7] indicated that the association between physician age and drinking persists even after stress was controlled.

Several lines of argument support the validity of our survey methodology.[6] Recent reviews of the validity of self-report alcohol consumption data have found that they are generally valid.[41–44] Eight studies of clinical populations (i.e., individuals with a drinking problem) revealed "a high degree of agreement between patients and collaterals (family members) with no consistent direction of error."[41] Five nonclinical population surveys "yielded somewhat similar findings."[41]

There is conflicting evidence regarding the possible biasing effect of survey nonresponse. One study of physicians[45] found that attaining a survey completion rate higher than that found here had little effect on the results. Another investigation[46] reported no systematic relationship between probability of response and amount of alcohol consumption. Two other studies (one on cigarette use[47] and another on alcohol use in Sweden[42]) found that heavy users were more likely than light users to be nonrespondents. The close match between our sample and the medical society's members, and the generally weak correlations between drinking measures and promptness of response to our survey by doctors, suggests that any bias in the present study is probably small.

While severe alcoholics who are no longer practicing medicine may have been omitted from the sampling list, research tells us that impaired physicians rarely drop out of medicine or do so only briefly.[48] During 1984, the state from which the

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**Gortmaker SL, Cheung L, Cleary P, Lampert S, Peterson N, Wechsler H: Results of the 1986 Harvard Health Survey at the Harvard School of Public Health, unpublished.

samples were drawn revoked the licenses of fewer than five physicians per 10,000 for all reasons.[49]

If the prevalence of alcohol use and abuse among physicians has been underestimated because of nonrespondents, that bias is probably no less true of pharmacists or medical students and pharmacy students. Since there is no reason to suspect *differential* bias, the results showing physicians and medical students to be no more vulnerable to alcoholism than pharmacists and pharmacy students are most likely valid. However, because of the nonresponse bias that has sometimes been found in surveys,[43,45,47] figures on the absolute prevalence of alcoholism for physicians and nonphysicians alike should be considered conservative.

Although the weight of evidence here and in previous studies[7,8,22,23,31,50,52] indicates that physicians and medical students are no more apt to abuse alcohol than are members of other professions, it is of some importance to the medical profession that, given the critical nature of their work, physicians are as vulnerable as their peers. The importance of the impaired physicians movement remains as an effective and humane model for handling problems of addiction. The findings reported here can best be used to remind us all that in any group where alcohol use is nearly universal and use rates are more than nominal, the risk of abuse and impairment cannot be ignored. ☐

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Comments Sought on Product Safety Aspects Examination

The Board of Certified Safety Professionals (BCSP) has offered the Product Safety Aspects Examination since October 1981. In that nine-year period, very few candidates for certification have selected this specialty examination. Currently, there are only 25 Certified Safety Professionals certified in Product Safety Aspects.

Maintaining a good quality examination is costly. The number of candidates taking the Product Safety Aspect Examination has been so small that the fee income from the candidates who take it does not cover the costs associatied with its development and maintenance.

Because of the limited interest and the cost of maintaining the examination, BCSP is considering whether or not to offer Product Safety Aspects as a specialty. In making this decision, the Board is attempting to determine if the demand for the examination will change in the future or if there are special circumstances that would justify its continutation even with a small number of candidates. Examples of factors which could result in increased demand for the examination are changing regulatory requirements or increasing employer/client emphasis on certification in Product Safety Aspects.

To assist in making this decision, the BCSP is soliciting comments and opinions from all interested parties. Comments on this issue may be submitted to: Michael K. Orn, CSP, Executive Director, Board of Certified Safety Professionals, 208 Burwash Avenue, Savoy, IL 61874-9510; Tel: (217) 359-9263; Fax: (217) 359-0055.

Effects on Serum Lipids of Adding Instant Oats to Usual American Diets

ABSTRACT

This study was designed as a test of the serum lipid response and dietary adaptation to recommended daily inclusion of instant oats in an otherwise regular diet. Hypercholesterolemic adults were randomly assigned to a control or intervention group. Participants in the intervention group were given packages of instant oats and requested to eat two servings per day (approximately two ounces dry weight), substituting the oats for other carbohydrate foods in order to maintain baseline calorie intake and keep weight stable. Serum lipids were measured in blood collected by venipuncture at baseline, four weeks, and eight weeks. Baseline mean total cholesterol (TC) levels were 6.56 mmol/L and 6.39 mmol/L for intervention and control groups, respectively. After eight weeks, mean serum total cholesterol of the intervention group was lower by −0.40 mmol/L, and mean net difference in TC between the two groups was 0.32 mmol/L (95% CI: 0.09, 0.54). Low-density lipoprotein-cholesterol was similarly reduced with mean net difference of 0.25 mmol/L (95% CI: 0.02, 0.48) between the two groups. Mean soluble fiber intake increased along with slight self-imposed reductions in mean total fat, saturated fat, and dietary cholesterol intake in the intervention group. Neither group changed mean body weight. Daily inclusion of two ounces of oats appeared to facilitate reduction of serum total cholesterol and LDL-C in these hyperlipidemic individuals. (*Am J Public Health* 1991; 81:183–188)

Linda Van Horn, PhD, RD, Alicia Moag-Stahlberg, MS, RD, Kiang Liu, PhD, Carol Ballew, PhD, Karen Ruth, MS, Richard Hughes, MD, and Jeremiah Stamler, MD

Introduction

Elevated serum cholesterol is an established risk factor for coronary heart disease (CHD).[1,2] About 1.5 million Americans have heart attacks each year. Evidence is extensive that lowering blood cholesterol reduces risk of heart attack, and slows, stops and even reverses atherogenesis.[2-7] The report of the National Cholesterol Education Program (NCEP) Adult Treatment Panel states that diet is the cornerstone of treatment for high blood cholesterol.[8] Epidemiological studies, controlled clinical trials, animal experimental data, and autopsy studies demonstrate that diet and plasma lipids are causally related.[9] Dietary fatty acids, dietary cholesterol, weight change, and water soluble fiber have all been observed to affect serum cholesterol.[10-12]

These and other studies have generated growing public interest in appropriate foods to help lower blood cholesterol. Physicians and other health professionals recommend reducing intake of total fat, saturated fatty acids, and cholesterol to maximize blood cholesterol lowering. A relatively simple step toward improving dietary behavior may be emphasis on daily inclusion of carbohydrate foods high in water soluble fiber, such as oats, to help displace some of the saturated fat and cholesterol containing foods. By advocating foods to eat rather than focusing solely on foods to restrict, dietary adherence and nutrient adequacy may be better achieved. To explore the efficacy and feasibility of this approach, this study was designed to measure the dietary changes and biochemical response of including moderate daily amounts of instant oats in the diet of free-living, hyperlipidemic men and women.

Methods

Participants

Men and women volunteers were recruited from a financial institution in Chicago. Capillary blood cholesterol levels were measured with use of the Reflotron* desk-top instrument, to confirm hypercholesterolemic status of individuals previously screened by the company and found to have serum cholesterol levels above 5.20 mmol/L. Eligibility criteria included: availability and willingness to follow study protocol, attendance at all data collection visits and completion of all food records, not on lipid-lowering drugs or anti-hypertensive medication, not diabetic, not pregnant or lactating, not greater than 1.5 times desirable body weight (1959 Metropolitan Life Insurance Weight Tables as standards), and not on a weight-loss diet. All participants signed informed consents. Participants were required to complete a three-day food record before the baseline data collection visit.

Data Collection

Individuals were randomized to one of two groups, stratified by sex and pre-

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From the Department of Community Health and Preventive Medicine, Northwestern University Medical School. Address reprint requests to Linda Van Horn, PhD, RD, Associate Professor, Department of Community Health and Preventive Medicine, Northwestern University Medical School, 680 North Lake Shore Drive, Suite 1102, Chicago, IL 60611-4402. This paper, submitted to the Journal December 14, 1989, was revised and accepted for publication June 22, 1990.

Editor's Note: See also related editorial p 159 this issue.

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*Boehringer-Mannheim, Indiana.

TABLE 1—Baseline Means and Standard Deviations for Group 1 and Group 2

	Group 1 Intervention	Group 2 Control
	n = 42	n = 38
Male/Female	21/21	19/19
Age (years)	42.9 ± 13.5	42.1 ± 11.9
Age Range (years)	25.0–76.2	22.2–66.9
Serum Total Cholesterol mmol/L	6.56 ± 0.82	6.39 ± 0.88
Low Density Lipoprotein Cholesterol mmol/L	4.58 ± 0.81	4.61 ± 0.73
Body Weight (kg)	76.2 ± 15.1	75.7 ± 13.4
Height (cm)	169.9 ± 10.8	169.8 ± 8.9
Body Mass Index (BMI)(kg/m^2)	26.2 ± 3.4	26.2 ± 3.8
Blood Pressure (mm Hg)	128/80 ± 14/10	129/79 ± 20/8

screen cholesterol level, above or below 6.34 mmol/L, prior to baseline visit. Group 1 was the treatment group and was asked to consume two packets (56.7 grams dry weight) of instant oats per day. This represented 5.6 grams of total dietary fiber per day, of which 3.4 grams was insoluble fiber and 2.2 grams was soluble fiber. Group 2 served as the control group and was asked to maintain usual intake throughout the study. Food records, height, weight, blood pressure, and serum blood lipids were measured at baseline and two subsequent visits, after four and eight weeks. Biochemical variables assessed were serum total cholesterol, triglycerides, low density lipoprotein cholesterol (LDL-C), high density lipoprotein cholesterol (HDL-C), and very low density lipoprotein cholesterol (VLDL-C). Total cholesterol, HDL-C, and triglycerides were measured enzymatically (Technicon Autoanalyzer)** VLDL-C and LDL-C were calculated from the other lipids.[13] Body mass index (BMI) was calculated as weight (without shoes), measured in kilograms on a beam balance scale, divided by height in meters squared (kg/m^2). Blood pressure was measured twice by trained, certified technicians using a standard sphygmomanometer.

To avoid weight change, the intervention group was briefly instructed how isocalorically to substitute oats for other carbohydrate foods in their diet. Choice of preparation methods, time of day to consume the oats, and whether the packets of oatmeal should be eaten together or at separate times were left to the participants. No other dietary changes were advised. The intervention and control groups met separately.

Analyses

A two-tailed, two sample t-test was used for the analyses. The computer software program SPSS[14] was used.

Split samples of blood were analyzed for 10 percent of participants for quality control. Technical error was calculated by the formula:

$$TE = \sqrt{\sum_{i=1}^{N} d_i^2/2N},$$

where d_i is the difference between the ith pair of values for a person and N is the number of pairs; technical error divided by group mean × 100 was used to estimate percentage error. Food records were edited and coded by trained, certified nutritionists at the Nutrition Coordinating Center in Minneapolis. The National Heart, Lung and Blood Institute (NHLBI) nutrient data base (tape no. 16) was used for the nutrient analyses.[15] Blinded duplicates of 10 percent of all food records were sent for recoding and analysis for external quality control.

Subgroup analyses were performed comparing observed versus predicted cholesterol responses between groups based on change in Keys Scores. The Keys Score: $1.35 \times (2\,\Delta S - \Delta P) + 1.5\,\Delta Z$—where ΔS stands for the difference between the two diets in calories for saturated fatty acids (SFA), ΔP is the percent of calories for polyunsaturated fatty acids (PUFA), and ΔZ is the difference in the square root of dietary cholesterol in milligrams per 1,000 calories per day—can be used to predict group change in serum cholesterol level.[16] Analysis of covariance was used to measure mean changes in serum cholesterol adjusted for study group, sex, body mass index, and baseline cholesterol.[6]

Results

Prescreen total cholesterol levels were measured on 245 individuals. Of those, 111 met eligibility criteria, signed consents, and were randomized prior to the baseline visit. Failure to attend the baseline visit with a completed food record constituted a drop out. In all, 80 individuals completed all data collection visits; 14 were dropped from the Intervention Group and 17 from the Control Group. The majority of dropouts (18) did not bring food records and/or did not attend the baseline visit; other reasons: ill (5), wanted to lose weight (6), pregnant (1), and "oat product was too sweet" (1).

Mean age for the total cohort was 42 years (Table 1). There were no differences between the two groups in the characteristics measured including mean lipid levels.

As presented in Tables 2 and 3, after four weeks of intervention, Group 1 mean cholesterol was lower by 0.28 mmol/L and mean net difference, or the difference of the difference, between groups was 0.20 mmol/L. After eight weeks, mean net difference in total cholesterol was 0.32 mmol/L (95% CI = 0.09, 0.54) and in LDL-C was 0.25 mmol/L (95% CI = 0.02, 0.48). There were no significant differences between the groups in HDL-C or triglycerides.

Technical errors for serum lipid measurements were: total cholesterol, 1.6 percent, HDL-C, 2.7 percent; and triglycerides, 3.2 percent.

Reported dietary intake data are summarized in Table 4. At visit 3, Group 1 reported slightly decreased total fat, dietary cholesterol, and saturated fat. For the control group, there were slight increases in these dietary values. These opposite changes across groups resulted in an absolute difference in Keys Score of 5.31 (p = 0.039) from visit 3 to visit 1. Intervention group food records indicated increased intake of total, soluble, and insoluble fiber, iron, calcium, vitamin B_6, folacin, niacin, magnesium, and retinol. Group 1 reported consuming an average of 1.9 packets (54 grams) of instant oatmeal each day. This represented 5.3 grams of total dietary fiber of which 3.2 grams was insoluble fiber and 2.1 grams was soluble fiber. There were no reports of side effects from eating the oats.

**Technicon, New York.

TABLE 2—Group Means and Standard Deviations at Visits 1, 2, and 3 for Serum Lipids, Lipoproteins and Blood Pressure

Variables	Visit 1 Baseline	Visit 2 4 Weeks	Visit 3 8 Weeks
Total Cholesterol			
Group 1 mmol/L	6.56 ± 0.82	6.28 ± 0.92	6.15 ± 0.86
Group 2 mmol/L	6.39 ± 0.78	6.32 ± 0.84	6.30 ± 0.82
LDL-C			
Group 1 mmol/L	4.58 ± 0.81	4.24 ± 0.87	4.16 ± 0.83
Group 2 mmol/L	4.61 ± 0.73	4.39 ± 0.75	4.44 ± 0.77
HDL-C			
Group 1 mmol/L	1.29 ± 0.34	1.36 ± 0.28	1.28 ± 0.29
Group 2 mmol/L	1.26 ± 0.32	1.35 ± 0.26	1.30 ± 0.25
VLDL-C			
Group 1 mmol/L	0.70 ± 0.50	0.68 ± 0.42	0.71 ± 0.48
Group 2 mmol/L	0.51 ± 0.23	0.58 ± 0.28	0.56 ± 0.24
Triglycerides			
Group 1 mmol/L	1.52 ± 1.09	1.48 ± 0.92	1.54 ± 1.04
Group 2 mmol/L	1.12 ± 0.51	1.26 ± 0.61	1.22 ± 0.53
Blood Pressure (mm Hg)			
Group 1	127.9/80.3 (± 13/10)	121.8/80.8 (± 11/8)	123.5/77.0 (± 14/11)
Group 2	129.1/79.1 (± 20/8)	124.8/80.0 (± 11/8)	124.2/77.2 (± 13/9)

Group 1 = Intervention (oats), n = 42.
Group 2 = Control, n = 38.

Mean group difference in body mass index from visit 1 to visit 3 was not significant (Table 4). Average weight change from visit 1 to visit 3 was less than one pound with a slight decrease in Group 1 and increase in Group 2. Blood pressure changes also were not significant (Table 2).

Subgroup analyses explored possible differences in cholesterol response according to baseline median levels of serum total cholesterol, BMI, Keys Score, and sex (data not shown). Greater reductions in total cholesterol were observed among those individuals above the baseline median serum total cholesterol level of 6.34 mmol/L and below the median BMI of 26.17. Men and women in the intervention group experienced similar reductions in serum total cholesterol. In regression analysis to adjust for baseline total cholesterol, sex, BMI, and study group (Table 5), serum total cholesterol at visit 3 was significantly different between the two groups. Changes in serum cholesterol from visit 1 to visit 3 were also significantly different between the two groups, independent of sex, changes in BMI, and the baseline serum cholesterol.

Discussion

In this study, consumption of a moderate amount of instant oats in the daily food pattern fostered blood cholesterol lowering in hyperlipidemic individuals. The data suggest that dietary adaptation to the practice of regularly eating oats potentiates favorable effects on blood total and LDL-cholesterol levels as well as on nutritional status by reducing intake of total fat, saturated fat, and cholesterol without weight gain and improving intake of commonly deficient nutrients, i.e., iron, calcium and vitamin B_6. Participants appear to have, at least in part, substituted oats for other carbohydrate foods isocalorically, rather than simply adding oats to the baseline intake. Proposed mechanisms for the effect of oats relate to soluble fiber's capacity to bind bile acids and interfere with fat absorption, to increase fecal excretion of bile acids affecting hepatic secretion of lipoproteins and to ferment colonic bacteria, thus forming gases and short chain fatty acids and affecting hepatic cholesterol synthesis.[11,17]

Our group has previously reported that in a free living, normolipidemic population, 2 oz. (56 grams) of oatmeal or oat bran led to additional decreases in serum cholesterol of 4.2 to 5.3 mg/dl (adjusted for the reduction in the control group) over what was attributed to dietary fat and cholesterol modification.[18] This finding was confirmed by a follow-up study where the addition of 2 oz. of oatmeal decreased serum cholesterol 4.7 mg/dl more than what was achieved by a fat-modified diet alone.[19] Subgroup analysis further showed greater reductions in participants with higher baseline serum cholesterol levels.[19] Other intervention studies in free-living, hypercholesterolemic individuals have shown that one cup of oat bran plus five oat bran muffins daily (100 grams oat bran) induced a significant decrease in plasma cholesterol.[20]

A review of intervention studies in this area noted that addition of soluble fiber from oats, oat bran, legumes, pectin, guar, or locust bean gums has been reported to reduce plasma total cholesterol by 6–19 percent.[10] Particularly in hypercholesterolemic individuals, addition of oats has been reported to induce total cholesterol and LDL-C reduction.[21,22] The combination of a fat modified diet plus regular intake of oat bran by hyperlipidemic individuals was also shown to be as effective in cholesterol lowering as cholestyramine resin or colestipol with fewer side effects and less cost.[23]

More recently, Swain, *et al,*[24] reported blood cholesterol reductions of 7 to 8 percent in free-living, normolipidemic individuals (baseline mean serum cholesterol of 4.8 mmol/l) who consumed 100 gram supplements of oat bran or wheat cereal over six weeks in a crossover design study. The supplements were consumed in the form of muffins or entrees representing 35.6 percent (oat bran) and 40.8 percent (wheat cereal) of mean caloric intake reported during the two study periods. The authors concluded that isocaloric substitution of either oat bran or wheat for other carbohydrates produced similar blood cholesterol lowering effects in these normolipidemic individuals.[24] It was further suggested that previous studies including similar quantities (100 grams per day or more) of oat bran may have simultaneously induced quantitative and qualitative shifts in fat intake through replacement with oat bran and that these shifts were more directly responsible for

TABLE 3—Mean Group^{a} Differences ± Standard Deviations for Serum Lipids

	V2 − V1 (Baseline—4 weeks)				V3 − V1 (Baseline—8 weeks)			
Serum Lipid	Group 1	Group 2	Net Diff V2 − V1	95% CI	Group 1	Group 2	Net Diff V3 − V1	95% CI
Total cholesterol mmol/L	−0.28 ± 0.59	−0.07 ± 0.50	0.20	−0.04, 0.45	−0.40 ± 0.55	−0.09 ± 0.43	0.32	0.09, 0.54
LDL-C mmol/L	−0.33 ± 0.68	−0.22 ± 0.42	0.11	−0.14, 0.36	−0.42 ± 0.58	−0.17 ± 0.42	0.25	0.02, 0.48
HDL-C mmol/L	0.07 ± 0.16	0.09 ± 0.19	0.02	−0.06, 0.10	−0.00 ± 0.18	0.04 ± 0.20	0.04	−0.04, 0.13
Triglycerides mmol/L	−0.04 ± 0.56	0.13 ± 0.29	0.17	−0.03, 0.37	0.02 ± 0.52	0.09 ± 0.37	0.07	−0.13, 0.27

^{a}Group 1 = Intervention (oats), n = 42,
Group 2 = Control, n = 38.
*p <0.05
**p <0.01

TABLE 4—Baseline^{a} Group^{b} Means, Standard Deviations and Visit 3 Differences in Nutrient Intake from 3-Day Food Records

	Baseline Means		Difference V3 − V1		
Dietary Factor	Group 1	Group 2	Group 1	Group 2	Difference Group 1 − Group 2
Calories	1909.5 ± 465.0	2007.4 ± 519.0	76.2 ± 537.0	13.9 ± 560.5	62.4
Protein (%kcal)	17.0 ± 3.0	16.0 ± 2.7	−0.90 ± 4.0	−0.03 ± 3.5	−0.9
Carbohydrate (%kcal)	44.6 ± 7.1	44.6 ± 8.2	2.7 ± 7.0	−1.4 ± 7.6	4.00*
Total Fat (%kcal)	36.8 ± 6.6	37.0 ± 6.9	−0.20 ± 7.0	2.8 ± 6.4	−2.9*
SFA (%kcal)	12.6 ± 2.6	12.6 ± 3.0	−0.50 ± 3.0	0.93 ± 3.2	−1.4*
MFA (%kcal)	13.8 ± 2.9	13.6 ± 2.9	0.04 ± 3.3	1.6 ± 2.9	−1.6*
PFA (%kcal)	7.4 ± 2.4	7.8 ± 2.4	0.27 ± 2.7	0.12 ± 2.6	0.2
P:S Ratio	0.61 ± 0.2	0.66 ± 0.3	0.08 ± 0.30	−0.04 ± 0.3	0.1
Keys Score^{c}	41.5 ± 9.3	41.1 ± 10.4	−2.4 ± 10.8	3.1 ± 12.3	−5.3*
Sodium (mg)	3235 ± 777	3371 ± 930	284 ± 1085	154 ± 1150	130.0
Cholesterol (mg)	264.8 ± 116.3	284.7 ± 131.5	−8.6 ± 158.9	23.9 ± 176.6	−32.5
Total Fiber (gm)	14.5 ± 5.2	16.0 ± 7.1	3.5 ± 4.8	−0.78 ± 5.6	4.2***
Soluble Fiber (gm)	5.2 ± 1.9	5.5 ± 2.3	1.4 ± 1.8	−0.48 ± 1.9	1.8***
Insoluble Fiber (gm)	8.6 ± 3.5	9.3 ± 4.5	2.2 ± 3.4	−0.07 ± 3.9	2.3**
Alcohol (%kcal)	2.5 ± 4.3	3.4 ± 4.4	0.67 ± 3.3	−1.4 ± 3.7	2.1
BMI (Kg/m^{2})	26.2 ± 3.4	26.2 ± 3.8	−0.1 ± 0.56	0.1 ± 0.6	0.2
Thiamin (mg)	1.6 ± 0.7	2.5 ± 3.2	0.7 ± 1.6	−0.4 ± 4.3	1.0
Riboflavin (mg)	1.8 ± 0.9	2.4 ± 2.4	0.3 ± 1.5	0.5 ± 2.4	0.3
Niacin (mg)	21.7 ± 8.3	27.5 ± 22.3	4.9 ± 12.0	−5.1 ± 20.8	10.0**
Folic Acid (mcg)	276.2 ± 153.0	286.9 ± 145.3	167.1 ± 208.6	−17.1 ± 134.6	184.3***
Vitamin B_6 (mg)	1.9 ± 0.9	2.1 ± 1.3	0.6 ± 1.3	−0.3 ± 1.1	0.9**
Vitamin B_{12} (mcg)	6.5 ± 7.2	5.5 ± 3.4	−1.1 ± 7.4	1.9 ± 7.7	2.9
Pantothenic Acid (mcg)	5.2 ± 3.4	6.0 ± 4.8	0.02 ± 3.9	−1.1 ± 4.6	1.1
Calcium (mg)	764.4 ± 385.8	729.2 ± 280.7	172.9 ± 356.3	9.7 ± 250.2	163.2*
Iron (mg)	15.9 ± 8.7	14.7 ± 4.8	11.9 ± 12.7	1.0 ± 6.4	10.9***
Zinc (mg)	12.0 ± 4.5	12.1 ± 7.1	1.2 ± 7.8	0.9 ± 7.6	0.3
Magnesium (mg)	274.8 ± 87.5	291.4 ± 102.6	34.7 ± 96.5	−13.9 ± 107.7	48.6*

^{a}Visit 1 = baseline
^{b}Group 1 = Intervention (oats), n = 42
Group 2 = Control, n = 38
^{c}Keys Score = $1.35 \times (2\Delta S - \Delta P) + 1.5\Delta Z$
Where ΔS is the change in percentage of calories for SFA, ΔP is the change in percentage of calories for PUFA and z is square root of dietary cholesterol in milligrams per 1,000 calories per day.
*p ≤ .05
**p ≤ .01
***p ≤ .001

the observed serum cholesterol reduction than any effect of the oat bran itself. Demark-Wahnefried, *et al*,[25] similarly reported that both fat-modified and oat-bran enhanced diets lowered serum cholesterol by 10-17 percent among free-living, hyperlipidemic individuals (group mean baseline serum cholesterol level was 7.15 mmol/l). Relatively small sample sizes, differences in baseline blood cholesterol levels, differences in dietary assessment methodologies and incomplete nutrient data bases (lacking soluble versus insoluble fiber data) have made meaningful comparisons between such studies difficult.

TABLE 5—Regression Coefficients for Models of Visit 3 Total Serum Cholesterol (mg/dl) and Change in Total Serum Cholesterol (Visit 3 – Visit 1).

Dependent Variable = Visit 3 Total Cholesterol

Independent Variable	Coefficient	S.E. Coefficient
Group (0 = Control, 1 = Intervention)	−11.25	4.33
Total Serum Cholesterol Visit 1 (mg/dl)	.85	.07
BMI at Visit 3 (Kg/m^2)	.23	.65
Sex (0 = women, 1 = men)	−3.02	4.42

Dependent Variable = Change in Total Serum Cholesterol (Visit 3 – Visit 1)

Independent Variable	Coefficient	S.E. Coefficient
Group (0 = Control, 1 = Intervention)	−9.00	4.26
Total Serum Cholesterol Visit 1 (mg/dl)	−.16	.07
Change in BMI (Visit 3 – Visit 1) (Kg/m^2)	9.05	3.67
Sex (0 = women, 1 = men)	−2.18	4.14

From a public health perspective, perhaps the most important finding common among these studies is that the combined decreased fat and increased carbohydrate approach is practical, feasible, nutritious, and efficacious in lowering blood cholesterol among normolipidemic as well as hyperlipidemic individuals.

The study reported here measured self-imposed changes in dietary behavior and lipid response resulting from daily inclusion of instant oatmeal in the otherwise regular diet of hyperlipidemic men and women. Some of the earliest studies exploring the simple addition of fiber to a usual diet did not report specific dietary changes or displacement invoked when oats were substituted for other possibly high fat foods.[10,20,21,23] A more recent study incorporated large amounts of oats, constituting a major share of the caloric intakes thus limiting the participants capacity for other foods.[24] Documentation of the dietary changes associated with oat intervention serves to more accurately attribute blood cholesterol response to the relevant dietary factors. This study represents an initial attempt to document these changes. Results of nutrient analyses suggest that the recommendation to incorporate oats on a daily basis led to self-initiated changes in dietary fat intake, as well as increased intake of several essential nutrients. These results also illustrate the feasibility of eating moderate amounts of oats daily, without weight gain or side effects reported when greater amounts are consumed.[24] The differences between observed versus predicted blood cholesterol values in the intervention group suggest that the minor changes in dietary lipid intake alone did not completely explain the observed reductions. The instant oats, presumably by increasing soluble fiber intake, and the slight weight loss appear to have contributed to the overall serum cholesterol response. The relatively high drop out rate prior to baseline may have contributed to the imbalance in mean baseline blood cholesterol levels between groups that, although not statistically significant, may have somewhat confounded these results. Intensified intervention aimed at further decrease in intake of total fat, saturated fatty acids and dietary cholesterol, increased carbohydrate and reduction of excess body weight, would predictably further reduce serum cholesterol levels among these individuals.

More research is needed to clarify dose responses and evaluate long term effects of increased soluble fiber intake among hyperlipidemic and normolipidemic individuals. Potential differences in response to such regimens requires further exploration. ☐

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Calendar for APHA 1991 Policy Development Process

The following schedule has been established for the development of APHA public policy resolutions and position papers in 1991. Full details of the Association's procedures for this activity can be obtained from:

AMERICAN PUBLIC HEALTH ASSOCATION, Attention: Richard Carson, 1015 Fifteenth Street, NW, Washington, DC 20005; (202) 789-5623

March 22 All proposed resolutions and position papers are due at APHA headquarters. They should be sent to the attention of the Joint Policy Committee. Proposals submitted via fax must be followed by hard copy within two days.

May 6–7 The Joint Policy Committee (JPC) meets in Washington to make an initial group assessment on each proposed policy statement. (Between March 22 and May 6, members of four reference committees will review each proposed policy statement. In addition, appropriate section chairpersons, SPIGs, and selected individuals are asked to review related proposed policy statements.)

May 24 All proposed statements are returned to authors with an inital assessment by the JPC.

June 28 Revised policy statements, to be considered by the 1991 Governing Council, must be back at APHA headquarters.

September Proposed policy statements are mailed to all APHA members in *The Nations's Health.* Copies are mailed to members of the Governing Council.

November 10 Annual Meeting—"Late-Breakers"—Submission deadline 6:00 pm (Sunday), for "late-breaking events." The chairpersons of the public hearings will be instructed that only those policy statements which address issues that have arisen between March 22 and the Annual Meeting will be considered as late-breaking events.

November 11 Annual Meeting—Public Hearings are conducted by the Reference Committees.

November 12 Annual Meeting—Joint Policy Committee meets in executive session to develop final recommendations for presentation of proposed resolutions and position papers to the Governing Council on Wednesday.

November 13 Annual Meeting—Governing Council votes on policy.

Validation of a Surveillance Case Definition of Carpal Tunnel Syndrome

Jeffrey N. Katz, MD, MS, Martin G. Larson, ScD, Anne H. Fossel, and Matthew H. Liang, MD, MPH

ABSTRACT

The National Institute for Occupational Safety and Health (NIOSH) has proposed a surveillance case definition for work-related carpal tunnel syndrome (CTS). The case definition requires the presence of median nerve symptoms; one or more occupational risk factors; and objective evidence of CTS including one of three physical examination findings or nerve conduction tests diagnostic of CTS. We evaluated the performance of the NIOSH case definition, restricting our analysis to cases in which physical examination findings served as the objective criterion. Nerve conduction studies were used as the gold standard. Seventy-eight workers were studied; 38 percent had CTS. The NIOSH case definition had sensitivity of 0.67 (95% CI = 0.57, 0.77), specificity of 0.58 (95% CI = 0.47, 0.69), and positive and negative predictive values of 0.50 (95% CI = 0.39, 0.61) and 0.74 (95% CI = 0.64, 0.84), respectively. Overall 38 percent of subjects were classified incorrectly. In a sample with a prevalence of 15 percent, as might be encountered in high risk workplaces, the positive predictive value would be 0.22. In conclusion, when physical examination findings serve as the objective criterion the performance of the case definition is modest reflecting the limited diagnostic value of its component tests and indicating that effective screening for CTS awaits improved diagnostic techniques. (*Am J Public Health* 1991;81:189–193)

Introduction

Carpal tunnel syndrome is a major problem in occupational health, particularly in occupations requiring highly repetitive and/or forceful hand motions.[1] Up to 15 percent of workers in the highest risk industries are affected annually,[2] 100 times more than expected in an age- and sex-adjusted population.[3] Costs per worker may exceed $15,000, including medical and indemnity expenses.[2,4]

Accurate and efficient diagnosis of carpal tunnel syndrome is important for research and clinical practice. Nerve conduction testing is generally regarded as the gold standard for the diagnosis of carpal tunnel syndrome.[5] The sensitivity of nerve conduction tests has been estimated at over 80 percent when compared with a clinical diagnosis of carpal tunnel syndrome in a community based study[6] and over 90 percent when compared with relief following surgery in 292 patients undergoing carpal tunnel release.[7] Electrophysiologic tests require 30 to 60 minutes, expensive equipment and experienced operators; thus, they may be inappropriate for many occupational surveillance efforts. Thermography[8] and vibrometry[9] are noninvasive but costly and have not been evaluated rigorously.

The clinical history and physical examination are inexpensive and simple to administer but are of limited value in the diagnosis of carpal tunnel syndrome. For example, nocturnal pain has a specificity of 0.28 and diminished two-point discrimination has a specificity of 0.32.[10] The Tinel and Phalen signs have sensitivities of 25–75 percent and specificities of 47–90 percent.[10–14] These data suggest that screening instruments comprised solely of historical and physical findings may not accurately identify workers with carpal tunnel syndrome.

The National Institute for Occupational Safety and Health (NIOSH) recently approved a surveillance case definition for work-related carpal tunnel syndrome[15] (see Figure 1). The NIOSH surveillance case definition of carpal tunnel syndrome requires that workers satisfy three criteria: symptoms of median nerve involvement; a history of occupational risk factors such as repetitive or forceful hand tasks or use of vibrating tools; and either the presence of physical examination findings—including the Tinel or Phalen signs or decreased pinprick sensation in the median nerve distribution of the hand—or diagnostic nerve conduction studies. To date the validity of NIOSH case definition has not been tested. The performance of the case definition is of particular interest when physical examination findings are used as the objective evidence required to fulfill criterion B (Figure 1). The purpose of this study is to test the NIOSH case definition in a sample of symptomatic workers. Physical examination findings serve as the sole objective evidence of disease (criterion B) in this

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Address reprint requests to Jeffrey N. Katz, MD, Arthritis Center, Brigham and Women's Hospital, 75 Francis Street, Boston, MA 02115. He is also Instructor in Medicine, Harvard Medical School; Dr. Larson is Associate Professor of Biostatistics, Harvard School of Public Health; Ms. Fossel is Research Associate, Robert Brigham Multipurpose Arthritis Center, Brigham and Women's Hospital; Dr. Liang is Associate Professor of Medicine, Harvard Medical School. This paper, submitted to the Journal February 27, 1990, was revised and accepted for publication July 13, 1990.

Editor's Note: See also related editorial p 161 this issue.

FIGURE 1—NIOSH Surveillance Case Definition of Work-Related Carpal Tunnel Syndrome*

Criteria A, B, and C must be met:

A. Symptoms suggestive of carpal tunnel syndrome: paresthesia, hypoesthesia, pain or numbness affecting at least part of the median nerve distribution of the hand.

B. Objective findings consistent with carpal tunnel syndrome:

Either, (1) One or more of the following physical findings: Tinel sign, Phalen sign, or decreased or absent sensation to pin prick in the median nerve distribution of the hand.

or, (2) Electrodiagnostic findings of median nerve dysfunction across the carpal tunnel.

C. Evidence of work relatedness: One or more of the following: Frequent, repetitive or forceful hand work on affected side; sustained awkward hand position; use of vibrating tools; prolonged pressure over wrist or base of palm; temporal relationship of symptoms to work or association with carpal tunnel syndrome noted in co-workers.

*adapted from reference 15

study and nerve conduction tests are used as the gold standard for the diagnosis of CTS.

Methods

An evaluation of clinical tests for carpal tunnel syndrome was carried out from January through June 1988. Details of the study have been reported.[10] Briefly, consecutive patients over 18 years of age with pain or paresthesia in the upper extremity who were referred to the Brigham and Women's Hospital Neurophysiology Laboratory were studied. A subset of this sample who were working or disabled as a result of their upper extremity condition are the subject of this report. Prior to neurophysiologic testing patients were evaluated by one of four examiners (a rheumatologist and three research assistants) who administered a questionnaire and performed a standardized physical examination. Examiners were blind to the results of nerve conduction tests.

Clinical data included demographic information, co-morbid conditions, occupation, and self-report of whether patients were exposed in their work to repetitive hand motion, vibration, or forceful wrist flexion or extension. The subjects also completed a hand pain diagram[16] which depicts dorsal and palmar views of both hands. Patients were asked to shade in areas corresponding to the location of their symptoms. Diagrams were rated blindly as "classic carpal tunnel syndrome," "probable," "possible," or "unlikely."

Moving two point discrimination[17] was performed with electrocardiogram calipers with tips set at four millimeters apart. The Tinel sign was performed by dropping the square end of a reflex hammer on the distal wrist crease. Pain or paresthesia in at least one of the first three digits constituted a positive response.[18] The Phalen sign was performed with the dorsal aspects of both flexed wrists in apposition for 60 seconds, and was considered positive if pain or paresthesia occurred in one of the first three digits.[19]

The presence or absence of CTS was established by the neurophysiologists with nerve conduction testing. The core protocol consisted of bilateral ulnar and median motor and sensory nerve conduction tests and electromyography of the abductor pollicus brevis. (Some subjects refused bilateral testing and/or electromyography.) Additional tests such as palm-to-wrist latencies[6] were performed according to the clinical circumstances at the discretion of the neurophysiologists. The neurophysiologists were blind to the results of physical examinations performed by the study investigators. The diagnosis of carpal tunnel syndrome required at least one of the following: median motor latency greater than 4.0 m/sec; sensory latency greater than 3.7 m/sec; and/or sensory velocity less than 50 m/sec. Ulnar nerve conduction and electromyography were used to exclude polyneuropathy, radiculopathy, and other diagnoses.

Criterion A of the NIOSH case definition (Figure 1) corresponded to a possible, probable, or classic rating on the hand pain diagram.[16] Criterion B corresponded to a positive Tinel or Phalen sign or diminished two point discrimination. The nerve conduction criterion was not used because our goal was to evaluate the performance of the case definition specifically when physical examination data are used as the objective criterion. Criterion C was satisfied if patients reported exposure on their jobs either to vibration, repetitive pinching, grasping, or wrist flexion or extension.

The sensitivity, specificity, odds ratio, positive predictive value (proportion with a positive test result who had carpal tunnel syndrome) and negative predictive value (proportion with a negative test result who did not have carpal tunnel syndrome) were determined for each of the diagnostic tests and for the NIOSH case definition.[20] Statistical analyses were performed with the chi-squared statistic using the SAS statistical package[21] on an International Business Machines P/S 2 Model 60 computer. Confidence intervals for sensitivity and specificity estimates were performed assuming a binomial distribution[22] and confidence intervals for predictive values were calculated with an iterative proportional fitting algorithm assuming a hypergeometric distribution.

Results

Seventy-eight symptomatic subjects were studied. Sixty-six were working at the time of evaluation and 12 were temporarily disabled as a result of their upper extremity problems. Patient characteristics are presented in Table 1. The mean age was 43.4 years. Sixty-three percent of subjects were female and 77 percent were White. Ten percent of patients had diabetes mellitus, 6 percent had thyroid disorders, and 4 percent had rheumatoid arthritis. These conditions were well controlled at the time of evaluation. Ten percent (eight patients) had prior carpal tunnel release surgery. Four of these eight patients were found to have carpal tunnel syndrome; in all four the previously operated hand was involved. Sixty-nine percent of subjects reported that their work involved repetitive hand motion, vibration, or forceful wrist flexion or extension. Thirty-eight percent of subjects had carpal tunnel syndrome documented by nerve conduction; diagnoses in the other patients are listed in Table 1.

Table 2 depicts diagnostic performance features of the individual components of the NIOSH case definition and of the case definition itself. In univariate analyses the presence of symptoms in at least part of the median nerve distribution of the hand was highly sensitive for carpal tunnel syndrome (0.93 [95% CI: 0.87, 0.99]) but non-specific (0.25 [95% CI: 0.15, 0.35]). The presence of occupational risk factors was also sensitive (0.83 [95% CI: 0.75, 0.91]) but nonspecific (0.40 [95% CI:

TABLE 1—Baseline Clinical and Demographic Features (N = 78)

Feature	Number of Subjects	(%)
Age (years)		
Less than 40	28	(35.9)
40–54	36	(46.2)
Greater than 54	14	(17.9)
Female	49	(62.8)
Duration of symptoms*		
Less than 2 months	18	(24.0)
2–12 months	31	(41.3)
Greater than 12 months	26	(34.7)
Prior carpal tunnel release	8	(10.3)
Occupation		
Professional (e.g. lawyer, doctor, business person)	15	(19.2)
Clerical/data processing	16	(20.5)
Manual handling	24	(30.8)
Manager (e.g. of store)	15	(19.2)
Artisan	8	(10.3)
Electrophysiological diagnosis		
Normal	22	(28.2)
Carpal tunnel syndrome	30	(38.5)
Ulnar neuropathy	10	(12.8)
Cervical radiculopathy	9	(11.5)
Other†	7	(9.0)
Total	78	100

*Data missing on three patients.
†Other diagnoses include brachial plexus lesions, polyneuropathy, other entrapments.

0.29, 0.51]). The Phalen sign was similarly sensitive (0.73 [95% CI: 0.63, 0.83]) but non-specific (0.36 [95% CI: 0.25, 0.47]), while the Tinel sign was less sensitive (0.62 [95% CI: 0.51, 0.73]) but more specific (0.66 [95% CI: 0.56, 0.77]). Two point discrimination was the least sensitive (0.23 [95% CI: 0.14, 0.32]) and most specific (0.82 [95% CI: 0.74, 0.91]) test for carpal tunnel syndrome.

The individual criteria were combined according to the NIOSH surveillance case definition and compared to actual diagnoses confirmed by nerve conduction tests. The case definition had sensitivity of 0.67 (95% CI: 0.57, 0.77), specificity of 0.58 (95% CI: 0.47, 0.69), positive predictive value of 0.50 (95% CI: 0.39, 0.61) and negative predictive value of 0.74 (95% CI: 0.64, 0.84). Overall, 48 of 78 (62 percent) subjects were classified correctly by the surveillance case definition. Bayes' Theorem[20] was used to calculate the predictive values in a population with prevalence of 0.15, as might be found in high-risk industries. In such a population the positive predictive value of the case definition would be 0.22 and the negative predictive value would be 0.91.

Discussion

Carpal tunnel syndrome has been associated with a variety of occupations including meat packing,[2] aircraft and bearing manufacturing,[1,23] sewing,[24] grocery checking,[25] and many others. Epidemiologic studies relating carpal tunnel syndrome to specific occupations have been limited by the lack of uniform, valid criteria for case ascertainment. The diagnostic gold standard, nerve conduction testing, is not suitable for many large surveillance efforts. Furthermore, simple clinical tests are of limited diagnostic value.[10]

The proposed NIOSH surveillance case definition is an important step toward standardization of surveillance. The present study is the first validation of the case definition against verified cases of carpal tunnel syndrome. This study suggests that when physical examination findings serve as the objective criterion the case definition has important limitations and must be used carefully. In our sample of workers with upper extremity complaints the positive predictive value of the case definition was 0.50 (95% CI: 0.39, 0.61) and the negative predictive value was 0.74 (95% CI: 0.64, 0.84). These results indicate that half of workers satisfying the case definition did not have carpal tunnel syndrome while a fourth of subjects not satisfying the case definition did. Overall, 38 percent of subjects were classified incorrectly.

It is critical to note that predictive values vary with the prevalence of disease. Carpal tunnel syndrome was more prevalent in our sample (0.38) than in other reported groups of symptomatic workers.[26,27] In a hypothetical sample of workers with a prevalence of CTS of 15 percent the positive predictive value of the NIOSH case definition would be just 0.22.

In evaluating the usefulness of the case definition it is important to distinguish between epidemiological surveillance, which provides information about disease prevalence in target populations, and screening, which identifies individuals at risk of disease.[28] The NIOSH case definition is a surveillance instrument intended for "use by state health departments to facilitate standardized counting of cases"; it is not intended for clinical use.[15] In surveillance programs designed solely to identify high-risk workplaces, simplicity and low cost are paramount. False negative and false positive results may compromise estimates of risk in these settings but do not adversely affect individual workers.[28] Our data suggest that the NIOSH case definition would be suitable for occupational surveillance as it is simple, inexpensive and accurate in 62 percent of subjects in our sample.

In contrast, inaccurate results exact higher costs when tests are used to identify possible cases for further testing and treatment. Here the cost of false negative test results is withholding treatment from workers who might benefit. The costs of false positives include worker anxiety, work absenteeism,[29] and the expense of disproving the screening diagnosis (e.g. $200–500 for a physician visit and nerve conduction tests). The modest sensitivity (0.67) and specificity (0.58) of the NIOSH case definition suggest that it should be used carefully if at all in screening for affected individuals or in clinical practice.

Other combinations of tests performed in parallel or in series would produce different trade-offs between sensitivity and specificity.[30] For example, we have demonstrated that the combination of a probable or classic hand diagram *or* a positive Tinel sign has sensitivity of 0.82 (95% CI: 0.75, 0.89) and specificity of 0.50 (95% CI: 0.41, 0.59), while a probable or classic diagram *and* positive Tinel sign has sensitivity of 0.39 (95% CI: 0.30, 0.48) and specificity 0.89 (95% CI: 0.83, 0.95).[10] However, no individual test or combination of tests offers sensitivity and specificity over 0.70 (Table 2) reflecting inherent limits in the diagnostic value of clinical tests for carpal tunnel syndrome.[10]

This study has several limitations. Subjects were referred to a hospital labo-

TABLE 2—Diagnostic Performance Characteristics of the NIOSH Case Definition and Its Component Parts

						In Sample with 15% Prevalence	
	Sensitivity	Specificity	Positive Predictive Value	Negative Predictive Value	Odds Ratio (95% CI)	Positive Predictive Value	Negative Predictive Value
Median n. symptoms	0.93	0.25	0.44	0.86	4.67 (0.97, 22.57)	0.18	0.96
Occupational risk factors*	0.83	0.40	0.46	0.78	3.28 (1.07, 10.06)	0.21	0.93
Tinel sign	0.62	0.66	0.53	0.74	3.17 (1.21, 8.30)	0.24	0.91
Phalen sign	0.73	0.36	0.40	0.70	1.50 (0.5, 4.3)	0.17	0.88
Median sensory loss†	0.23	0.82	0.47	0.62	1.45 (0.46, 4.52)	0.19	0.86
Case definition	0.67	0.58	0.50	0.74	2.80 (1.08, 7.25)	0.22	0.91

*Exposure to repetitive pinching, grasphing, wrist flexion or vibration.
†Moving two point discrimination of 4 mm.

ratory for neurophysiologic evaluation. Some may have had particularly severe carpal tunnel syndrome, raising the sensitivity of diagnostic maneuvers, or atypical syndromes lowering specificity.[31] While all subjects were working or disabled it is not clear what proportion of cases were unequivocally work-associated or whether the case definition would perform differently in such cases. Also, we limited our analysis to cases in which criterion B was satisfied by physical examination rather than nerve conduction findings. Thus our analysis likely underestimates the specificity of the case definition. In addition, while nerve conduction tests are the best gold standard available, false negatives are well documented[7] and the false positive rate is not precisely defined, raising the possibility of misclassification error. Furthermore, occupational risk was estimated on the basis of self-reported exposure to vibration and repetitive finger and wrist flexion which is less accurate than detailed ergonomic analyses of job requirements. Finally, carpal tunnel syndrome is just one of a variety of occupation-associated upper extremity disorders. Symptomatic workers who do not have carpal tunnel syndrome may have other occupation-related cumulative trauma disorders[26] and might benefit from job modification and treatment.

We conclude that when the physical examination is used as the objective evidence to satisfy criterion B, the NIOSH case definition yields a significant number of false positive and false negative results. False positives will be especially common in groups of workers with a low prevalence of carpal tunnel syndrome. While the case definition appears to be suitable for epidemiologic surveillance, these inaccuracies make the case definition suboptimal for use in screening individual workers. The modest performance of the case definition reflects limitations inherent in its component tests, suggesting that accurate non-invasive testing for occupational carpal tunnel syndrome awaits the development and validation of better diagnostic technologies. ☐

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Free Information Packet on Women and AIDS

In response to the growing need for information on women and HIV (human immunodeficiency virus)/AIDS (acquired immunodeficiency syndrome), the National AIDS Information Clearinghouse has increased its efforts to collect and disseminate information for and about women with HIV. Results of this effort led to the development of a new information packet which provides resources offered by the Clearinghouse. The Women and AIDS Information Packet developed by the Clearinghouse contains:

- A fact sheet describing Clearinghouse resources and services for women: e.g. to locate funding opportunities for women's programs or to find child care and housing assistance programs.
- Customized, topical searches of Clearinghouse databases: a selected list of educational materials for professionals concerned about HIV and AIDS and women; a list of materials for use in HIV education activities with women; a list of organizations offering services for women.
- A publications list of HIV and AIDS eductional materials offered by the Clearinghouse including brochures, posters, and videotapes.
- *A Selected Guide to Information about Women and HIV/AIDS* for use by health professionals, graduate or undergraduate students, or other interested persons. The guide lists books, indexes, online information sources, newsletters, journals, directories, and organizations that provide information about women and AIDS.

To obtain the free information packet, contact National AIDS Information Clearinghouse, P.O. Box 6003, Rockville, MD 20850; Toll-free: 1-800-458-5231, TTY/TDD: 1-800-243-7012, FAX: 1-301-738-6616; International line: 1-301-217-0023; Hours: 9am–7pm, EST, M–F.

Public Health Briefs

Behavioral, Health and Psychosocial Factors and Risk for HIV Infection among Sexually Active Homosexual Men: The Multicenter AIDS Cohort Study

Lili Penkower, PhD, Mary Amanda Dew, PhD, Lawrence Kingsley, PhD, James T. Becker, PhD, Paul Satz, PhD, Frederick W. Schaerf, MD, PhD, and Kathleen Sheridan, JD, PhD

ABSTRACT

We examined whether 644 homosexual men who engaged in receptive anal intercourse were at particularly elevated risk for seroconversion if they also possessed specific behavioral, health or psychosocial vulnerability characteristics. Of 11 potential factors examined, heavy drinking, moderate to heavy drug use, and younger age were significantly related to seroconversion. These variables were also associated with an increased number of sexual partners, anonymous sex, and failure to use condoms. (*Am J Public Health* 1991; 81:194–196)

Introduction

Host vulnerability to illness is a complex process involving biologic, psychologic and social factors.[1-4] With regard to acquired immunodeficiency syndrome (AIDS), the etiologic agent, human immunodeficiency virus (HIV), is known and two behavioral factors, intravenous drug use and high-risk sexual behavior, transmit the virus.[5] Cofactors, increasing the likelihood that HIV-exposed individuals will become infected, have been suggested.[6-8] They include pre-existing health and behavioral factors (e.g., smoking, recreational drug use, alcohol consumption) and psychosocial factors (e.g., depression, lack of social support). We examined the impact of these variables on seroconversion risk in homosexual men engaged in high-risk sexual activity.

Methods

A sample of 1,045 bisexual and homosexual men was drawn from the total cohort of men enrolled in the Multicenter AIDS Cohort Study (MACS) between 1984–87. Participants in this four-site, longitudinal study were examined semi-annually (see Kaslow, *et al*,[9] for further cohort description). Our sample included all 209 men who: became HIV seropositive (HIV+) between 1984–87; and had tested seronegative (HIV−) six months before. Data from their last HIV− visit were utilized as "baseline" data; data from their first HIV+ visit were designated "follow-up" data. A comparison sample of 836 men, who were HIV− throughout 1984–87 and represented the four sites in the same proportion as the HIV+ sample, was randomly selected. For each HIV− man, we randomly selected two consecutive waves of interviews (six months apart) which we designated as the "baseline" and "follow-up" visits.

High-risk sexual behavior was defined as receptive anal intercourse in the prior six months.[10-12] At each interview, additional variables in the previous six months were assessed: number of sexual partners with whom respondents had receptive anal intercourse; number of anonymous partners; number of partners with whom respondents had used condoms during anal intercourse; and number of partners who themselves used condoms during anal intercourse with respondents.

Three categories of vulnerability factors, potentially affecting seroconversion risk, were measured at baseline (see first two columns of Table 1 for descriptive statistics). Absolute numbers of CD4+ cells were counted by direct immunofluorescence with use of monoclonal antibodies with an EPICS-C flow cytometer.[13] All variables were available for all men.

Results

At baseline, high-risk sexual behavior was reported by 87.0 percent of the 209

Address reprint requests to Lili Penkower, PhD, Department of Clinical Epidemiology and Preventive Medicine, University of Pittsburgh School of Medicine, 3434 Fifth Avenue, Pittsburgh, PA 15260. Drs. Dew, Kingsley, and Becker are also the University of Pittsburgh; Dr. Satz is with the University of California at Los Angeles; Dr. Schaerf is with Johns Hopkins Medical Institutions; and Dr. Sheridan is with Northwestern University Medical School. This paper, submitted to the Journal June 1, 1990, was revised and accepted for publication September 27, 1990.

TABLE 1—Relationship of Pre-existing Vulnerability and Subsequent Serostatus among High-risk Sexually Active Men (N = 644)

	Seropositive (N = 181)		Seronegative (N = 463)		Regression[h]	
Vulnerability Factor	n	%	n	%	Odds Ratio	95% Confidence Interval
Behavior						
Alcohol consumption (% heavy)[a]	62	34.3	86	18.6	1.39	1.12, 1.72
Smoking (% yes)	78	43.1	143	30.9	1.22	.99, 1.50
Recreation drug use (% moderate-heavy)[b]	92	50.8	166	35.9	1.25	1.03, 1.51
Immune Status and Health History						
CD4+ Cells (% ≤ 700)[c]	44	25.9	102	23.4	1.16	.93, 1.45
Chronic physical illness (% yes)[d]	8	4.4	14	3.1	1.17	.73, 1.97
Psychosocial and Background						
Depression (% highly symptomatic)[e]	35	19.7	79	17.1	.89	.68, 1.16
Self-esteem (% low)[f]	13	7.3	22	4.8	1.32	.88, 1.99
Confidant (% none)[g]	10	5.6	15	3.2	1.30	.83, 2.03
Age (% ≤ 35 years)	122	67.4	274	59.2	1.25	1.02, 1.52
Education (% ≤ high school)	23	12.8	41	8.9	1.13	.83, 1.53
Race (% non-White)	11	6.1	13	2.8	1.20	.73, 1.97

[a] Cahalan quantity-frequency index.[14]

[b] recreational drugs include at least weekly use of marijuana, cocaine, nitrite inhalants, barbiturates, or amphetamines.

[c] fell in lowest quartile of sample distribution.

[d] chronic illnesses include cancer, diabetes, kidney disease, lupus, vasculitis, rheumatoid arthritis, ankylosing spondylitis, Graves disease, Reiter's syndrome, or myasthenia gravis.

[e] scores of 16+ on the Center for Epidemiologic Studies Depression Scale (CES-D).[15]

[f] one item from Rosenberg Scale.[16]

[g] Brown & Harris[17]

[h] Goodness of Fit χ^2 (1) = 168, p = .011.

respondents who had seroconverted by follow-up, compared to 56.3 percent of the 836 seronegative men. The strong association between anal receptive sexual practices and seroconversion led us to focus on the 644 men who had engaged in high-risk sexual behavior at baseline.

After controlling for all other vulnerability factors, baseline heavy alcohol consumption, moderate to heavy drug use, younger age and, to a lesser extent, smoking were associated with seroconversion (Table 1).

We found no evidence that heavy drinkers, moderate to heavy drug users or younger men had fewer absolute numbers of CD4+ cells at baseline. Regarding the high-risk behaviors at baseline, however, heavy drinkers were indeed more likely to have engaged in receptive anal intercourse with more partners, had more anonymous sex, and were less likely to have either used condoms themselves or had partners who used condoms during anal intercourse. Similarly, moderate to heavy drug users reported more sexual partners and more anonymous sex. Younger men were more likely to have engaged in high-risk sex with more partners (data not shown; available on request to authors).

Discussion

The MACS' longitudinal design provided an unusual opportunity to examine data on behavioral, health, and psychosocial factors assessed prior to evidence of HIV seroconversion. Among men already at high risk for HIV infection, heavy alcohol consumption, moderate to heavy drug use and younger age further increased their likelihood of seroconversion by the follow-up visit.

Of these three factors, heavy alcohol consumption was most strongly associated with subsequent seroconversion. This finding is consistent with other evidence that alcoholics are at greater risk for infections of many types, and also have elevated morbidity and mortality rates.[18,19] Molgaard, *et al*,[20] have suggested that the importance of alcohol may derive from an association with impaired immunity and/or its behavioral effect on risky sexual practices. Our data suggest that its link to sexual practices may be of relatively greater concern. Thus, consistent with Kaslow, *et al*,[21] we found that CD4+ cells (albeit a limited measure of general immune function) bore little relation to heavy drinking. However, among men who all reported engaging in receptive anal intercourse heavy drinkers further elevated their risk of HIV exposure by engaging in sex with more partners, having more anonymous sex and by failing to employ condoms during high-risk sex.

Results for drug use and younger age were similar to alcohol. While drug users did not have lower CD4+ cell counts, they were more likely to engage in anonymous sex and had more sexual partners. It is noteworthy that, in comparison to all other drugs we examined, nitrite inhalant (popper) use was most strongly associated with seropositivity risk (odds = 11.83, 95% CI = 1.34, 3.04). Moss, *et al*,[22] recently reported a similar association. While no differences were observed in CD4+ cells among younger men, they were more likely to engage in high-risk sex with more partners.

We recognize several limitations in our study. First, a restricted number of potential risk factors were studied; many other factors (e.g., additional immune parameters, anxiety, nutrition, specific types of social support) remain to be investigated. Second, two of the psychosocial variables, self-esteem and social support, were each determined by a single question, and may have been inadequate measures of these constructs.

Nevertheless, our results suggest that efforts at controlling the spread of HIV infection may be improved by identifying individuals in high-risk groups whose other characteristics further heighten their already elevated likelihood of becoming infected. From a public health standpoint, it is imperative that the risks associated with substance abuse be recognized and that programs to educate and treat these addictions be made available and accessible. This study demonstrates empirically that individuals who engage in risky sexual practices and who are also heavy drinkers and/or drug users are at increased risk of becoming infected with HIV.

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Responsible Alcohol Service: A Study of Server, Manager, and Environmental Impact

Beth Howard-Pitney, PhD, Michael D. Johnson, PhD, David G. Altman, PhD, Rodney Hopkins, MS, and Noreen Hammond

ABSTRACT

A responsible alcohol-service training program was evaluated for its impact on changing beliefs, knowledge, and behavior in 97 servers and 43 managers and on changing establishment policies that encourage safer drinking environments. The training program had a significant impact on changing the beliefs and knowledge of both servers and managers. Observation 4 to 6 weeks after training showed no effects on server behavior, but there was a tendency toward more establishment policies compared with controls. (*Am J Public Health* 1991;81:197–199)

Introduction

Deaths and injuries associated with driving while under the influence of alcohol are important public health problems.[1] In 1988, half of all traffic fatalities were alcohol related.[2] Making drinking establishments the site of interventions to reduce alcohol-related accidents has increased in popularity. Establishments may be held legally responsible for the behavior of customers who leave under the influence of alcohol[3–5]; and a growing number of states and cities require that all servers of alcohol receive training in responsible alcohol service.

Evaluations of several alcohol-service training programs suggest that they can have some impact on the drinking behavior of patrons.[6–9] While some programs focus primarily on teaching servers how to intervene with intoxicated customers, others seek to be more comprehensive by focusing on servers, managers, and environmental features of the establishment.[10–11] The Responsible Beverage Service (RBS) Training Program studied in the current research focused on all three of these elements.[12]

Methods

A newly expanded Utah law required that servers of alcohol and managers of establishments that serve alcohol participate in a state-certified training program every 3 years. Ninety-seven servers from 19 restaurants and bars and 43 managers from 21 establishments (26 different establishments in all) in Park City, Utah, attended a 1-day training session. Fourteen establishments matched on type and size were randomly selected from the remaining city establishments to serve as controls. They had not been exposed to this training, which was new in this community.

Servers and managers attended separate programs that taught the physical and behavioral effects of alcohol and strategies for providing more responsible alcohol service using a combination of lecture, role play, and feedback. Managers also received instruction on developing company policies to change the drinking environment. Participants completed pre- and post-training questionnaires measuring beliefs and knowledge about responsible alcohol-service practices. Further details are available on request.

One month following the RBS training, trained pairs of observers (each consisting of a man and a woman) with a mean age of 31 assessed server behavior and current policies in treatment and control establishments. The observers were blind to the treatment conditions, and the servers were unaware that they were being observed.

Observers first recorded ID-checking practices for the first three patrons both observers agreed looked 30 years old or younger (a practice taught to servers in the training) and environmental signs of responsible alcohol service in one establishment (described later). At a second establishment, one member of the pair drank three alcoholic beverages and tried to order a fourth in 1 hour. This amount ensured that the drinking observer was well on the way to being legally drunk (0.08 percent blood alcohol level in Utah). The consumption was such that if they followed their training, servers should have been able to interpret this drinking behavior as appropriate for intervention. The nondrinking observer recorded servers' interventions.

Results

T-tests on the pre- and post-test measures indicated that servers and managers

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Address correspondence and reprint requests to Beth Howard-Pitney, Stanford Center for Research in Disease Prevention, Stanford University School of Medicine, 1000 Welch Road, Palo Alto, CA 94304-1885. Drs. Howard-Pitney, Johnson, and Altman are with the Center for Research in Disease Prevention at Stanford University. Mr. Hopkins and Ms. Hammond are with Mountain Community Health Choices in Park City, Utah. This paper, submitted to the Journal December 20, 1989, was revised and accepted for publication June 21, 1990.

TABLE 1—Mean Number of Interventions by Servers in Treatment and Control Establishments

Type of Intervention	Treatment Servers (n = 22)	Control Servers (n = 19)
Communicated that server was aware of customer's drinking behavior	0.14	0.32
Slowed the delivery of drinks to the table	0.18	0.26
Suggested customer order a nonalcoholic drink or food	0.23	0.37
Offered to find transportation for an intoxicated customer	0.00	0.00
Explained that house policy does not allow the server to serve another alcoholic drink to customer	0.18	0.16
Verified that the nondrinking customer would do the driving	0.00	0.00
Refused to serve customer another alcoholic drink	0.23	0.16
Notified management if intoxicated customer became problematic	0.00	0.00
Mean of all interventions	0.95	1.26

Note: Data are based on observations by a nondrinking observer of server interventions with the drinking observation partner.

TABLE 2—Percent of Treatment and Control Establishments Showing Responsible Alcohol-Service Practices

Type of Practice Observed	Treatment Establishment (n = 13)	Control Establishment (n = 11)
Did not serve pitchers of beer or mixed drinks	46%	73%
Did not serve doubles	85%	82%
Did not offer happy hours	77%	82%
Had signs advertising low- or non-alcoholic drinks	38%	18%
Had transportation information	15%	9%
Had no employee drinking on the job	92%	73%
Mean of all policies	3.54	3.36

increased their knowledge and showed improvement in their beliefs that customers would respond favorably to responsible alcohol service and policies $p < .001$, all measures. The training was less successful in teaching servers and managers new strategies for intervening with customers, as reported in open-ended questions. (Data are available on request.)

Table 1 reports data on observations of interventions by servers. No differences were observed between treatment (n = 22) and control (n = 19) servers on any intervention nor on a summed average of eight possible interventions.

Observations of establishment practices were conducted in 13 of the 21 establishments with trained managers and 11 of the 14 control establishments (Table 2). Treatment establishments were more likely to have signs advertising low-alcohol or nonalcoholic drinks and to show less evidence of employees drinking on the job than control establishments, but they were also more likely to serve pitchers of beer or mixed drinks. These findings are tentative because of the small sample sizes. In addition, checking IDs of patrons under 30 was observed at only 14 treatment establishments (9 percent) and 14 control establishments (18 percent), a nonsignificant difference.

Discussion

Questionnaire data indicated that the training program had a positive impact on improving knowledge and beliefs. Observational data collected 4 to 6 weeks after training about the effectiveness of the training program were inconclusive given the small sample sizes and design limitations of the study. Diffusion of training from treatment to control conditions, while unlikely, cannot be completely ruled out; neither can preexisting differences between conditions, although matching on establishment type and size should have increased the comparability between conditions.

Overall, the observations seemed to indicate relatively little impact of the training program on server behavior but did show a tendency for the treatment establishments to engage in more responsible practices than control establishments. Although tentative, these findings suggest that other programs may want to incorporate training on policy issues in addition to server-intervention training.

Responsible alcohol-service programs should consider offering programs in the working environments of managers and servers to better convince staff of the benefits of responsible alcohol-service practices. On-site, tailored feedback on appropriate practices might provide more potent training and a better diagnosis of establishment features that create barriers to responsible alcohol service.[9]

Furthermore, one-on-one training with managers, including follow-up support, may be needed to encourage establishments to develop policies and to help them overcome obstacles to implementation. In the current study, some managers may not have had the authority or support to implement new policies. Previous research suggests that before responsible alcohol-service practices can become fully institutionalized, support at all levels must be obtained.[5,9]

State and local regulations also can do more to encourage responsible alcohol-service practices. States that mandate training could require documentation of changes in policy and server behavior to encourage implementation of responsible practices. Regulations banning practices such as happy hours and serving doubles should be combined with educational strategies to convince establishments that practicing responsible alcohol service will not put them at an unfair competitive advantage or adversely affect customer volume and tips. Future research on the role drinking establishments can play in reducing alcohol-related problems should continue to focus on comprehensive strategies including government regulation, establishment policies, and server behavior. ☐

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Increased Risk of Ectopic Pregnancy with Maternal Cigarette Smoking

ABSTRACT

As part of a case-control study of ectopic pregnancy, we evaluated the potential etiologic role of cigarette smoking. Maternal cigarette smoking at the time of conception was associated with an increased risk of ectopic pregnancy with a dose-response relationship (adjusted odds ratios: 1.30 to 2.49). On the other hand, partner's smoking was not associated with ectopic pregnancy. The study provides a supplementary argument towards a causal effect of smoking in the development of ectopic pregnancy. (*Am J Public Health* 1991;81:199–201)

Joël Coste, MD, Nadine Job-Spira, MD, Hervé Fernandez, MD

Introduction

During the past two decades, the incidence of ectopic pregnancy has doubled or tripled in many parts of the world.[1] It currently constitutes 1.2 to 1.4 percent of all reported pregnancies[1], and remains the leading cause of maternal death during the first trimester of pregnancy in industrial countries.[2] The main identified risk factors are pelvic inflammatory disease (PID) and sexually transmitted diseases (STDs), previous ectopic pregnancies, pelvic surgery, previous use of intrauterine device (IUD), and oral contraceptives and IUD use at the time of conception.[1] Several studies have shown a positive association with cigarette smoking.[3–8] In most of them, smoking was not the primary hypothesis and the precise quantity of cigarette smoked was not investigated. Furthermore, some important confounding factors (e.g., STDs) were not taken into account.

As part of a case-control study of ectopic pregnancy, we further investigated the potential etiologic relationship between cigarette smoking, correcting for the above deficiencies wherever possible.

Methods

A case-control study was conducted during 1988 in seven large maternity hospitals in the Paris area (France). The cases were women ages 15–44 years, whose diagnosis had been confirmed by coelioscopy or laparotomy (n = 279). For each case, the first woman who delivered in the same center following the operation of the index-case was eligible as a control. When the woman selected as a control refused to be interviewed (n = 7), the woman immediately following who delivered in the same center was taken instead.

The interview collected standardized information on reproductive history, birth control practices history, sexual history, medical history including STDs and PID,

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Address reprint requests to Nadine Job-Spira, MD, INSERM U. 292-Hôpital de Bicêtre, 78, rue du Général Leclerc, 94275 Le Kremlin-Bicetre Cedex, France. Dr. Coste is also with that hospital unit; Dr. Fernandez is with INSERM U. 187, Maternité Antoine Béclère, Clamart Cedex France. This paper, submitted to the Journal February 28, 1990, was revised and accepted for publication June 12, 1990.

TABLE 1—Risk of Ectopic Pregnancy Associated with Sociodemographic Characteristics

Variables	Controls (n = 279) Number/ Percent of exposed		Cases (n = 279) Number/ Percent of exposed		Crude Odds ratio (95% CI)
Maternal Age (years)					
<20	10	3.6	4	1.4	0.63 (0.18, 2.18)
20–24	47	16.8	30	10.8	1
25–29	114	40.9	80	28.7	1.10 (0.64, 1.89)
30–34	72	25.8	91	32.6	1.98 (1.14, 3.44)
35–40	28	10.0	55	19.7	3.08 (1.61, 5.87)
≥40	8	2.9	19	6.8	3.72 (1.45, 9.57)
Nationality					
French	188	67.4	183	65.6	1
Foreign	91	32.6	96	34.4	1.08 (0.76, 1.53)
Educational Level					
None	13	4.7	11	4.0	0.77 (0.34, 1.78)
Primary	45	16.2	44	15.8	0.89 (0.56, 1.43)
Secondary	151	53.9	165	59.0	1
Higher	70	25.2	59	21.2	0.77 (0.51, 1.16)
Single Woman					
No	254	91.0	253	90.7	1
Yes	25	9.0	26	9.3	1.04 (0.59, 1.86)

surgical history, conditions of conception and sociodemographic characteristics. Maternal cigarette smoking was assessed by the number of cigarettes smoked per day at the time of conception. Information about the main sexual partner included: number of cigarettes smoked per day, educational level, socio-occupational class, and number of STDs during the past six months. Odds ratios (OR) and 95% confidence interval (CI) were used to describe the association between ectopic pregnancy and smoking.[9] Unconditional multiple logistic regression was performed to control for maternity hospital and to adjust for the confounding effects of other exposures.[10] The proportion of cases that can be attributed to cigarette smoking was assessed by the adjusted population attributable fraction calculated with a formula given by Walter,[11] and adapted for multivariate settings by Bruzzi, *et al.*[12]

Results

The main sociodemographic characteristics of cases and controls were similar except for age (Table 1).

A higher proportion of cases (40.1 percent) than controls (29.7 percent) were smokers at the time of conception (OR = 1.64, 95% CI: 1.15, 2.33). Furthermore, when smoking was analyzed as a categorical variable, we observed a clear dose gradient relationship between cigarette smoking and ectopic pregnancy (Table 2) (OR = 1.21 to 1.64, trend tendency $p < 0.05$). Variables adjusted for in the logistic regression model included: age, appendectomy, prior ectopic pregnancy, prior tubal surgery, prior spontaneous abortion, previous use of IUD, PID, induced conception cycle, use of IUD, combined contraceptive pill or progestative micropill at the time of conception. Adjustment slightly increased the crude estimates (Table 2). On the other hand, there was no association with partner's smoking (OR = 0.82, 95% CI: 0.58, 1.16). Since the adjusted OR of the ectopic pregnancy association with cigarette smoking was 1.68 and the prevalence of smoking among the controls was 0.30, the adjusted population attributable fraction in our population was 17 percent.

Discussion

We found an increased risk of ectopic pregnancy in women who were smokers at the time of conception. This association is supported by the dose-response relationship observed even after adjustment for identified confounding variables. Conversely, no such association was found with partner's smoking, an association not previously investigated. These results seem to eliminate possible masked behavioral confounding factors and strengthen the argument that cigarette smoking was a true direct risk factor of ectopic pregnancy.

Several potential sources of bias must be considered. Since cases and controls had similar demographic characteristics and since adjustment for the center was performed, selection bias is probably limited. However, recall bias which may have led to an overestimation of the risk, bias associated with the misclassification of the recent quitters (probably small according to Chow),[7] and the confounding effects of unidentified risk factors cannot be totally excluded.

Our findings are consistent with six previous epidemiological studies which collected information on maternal smoking habits.[3-8] Despite methodological differences, these studies all found a positive association. Chow, *et al*,[7] found a positive association between smoking at the time of conception and ectopic pregnancy (adjusted OR = 2.2), but failed to demonstrate a dose-response relationship. In the study of Handler, *et al*[8], the estimated relative risk associated with smoking rose from 1.4 for a woman smoking fewer than 10 cigarettes per day to 5.0 for 30 or more cigarettes per day. In that study, however, smoking was assessed during pregnancy. Because the pregnancies of the controls were longer, these women had greater possibility than cases to stop smoking during pregnancy and this may have led to an overestimation of the relative risk.

Alterations of tubal motility and ciliary function following injections of nicotine have been demonstrated.[13] Nicotine also delays blastocyst implantation and alters spacing of implantation sites[14] in animals. The possible role of low estrogen-levels is advocated by some authors.[15] Several findings support this hypothesis: women smokers have lower estrogen-levels than controls[16] and tubal contractions are clearly under estrogenic control.[17] Another explanation for the deleterious effect of smoking involves the reduced immunity in cigarette smokers.[18,19] The altered immunity may affect the tubal response to inflammation, resulting in an increased frequency of PID.[7] However, the association between smoking and PID has not yet been extensively explored in epidemiological studies.

Our observed population attributable fraction of nearly 20 percent indicates the public health importance of this association. If a factor is a causal agent of the disease, the reduction in the rate which would occur if exposure is prevented is equal to this proportion.[11]

In conclusion, our study provides a supplementary argument for the existence of a causal link between cigarette smoking and ectopic pregnancy. Smoking is widespread among women in many countries. In some cases, the prevalence among young women is ever increasing.[20,21] The

TABLE 2—Risk of Ectopic Pregnancy Associated with Cigarette Smoking

Variables	Controls (n = 279) Number/Percent of exposed		Cases (n = 279) Number/Percent of exposed		Crude Odds ratio (95% CI)	Adjusted* Odds ratio (95% CI)
Current Smoking† (woman)						
No	194	69.5	160	57.3	1	1
Yes	83	29.7	112	40.1	1.64 (1.15, 2.33)	1.68 (1.11, 2.55)
Unknown	2	0.8	7	2.6		
Number of Cigarettes/Day§ (woman)						
0	194	69.5	160	57.3	1	1
1–10	40	14.3	40	14.3	1.21 (0.75, 1.97)	1.30 (0.73, 2.32)
11–20	37	13.3	60	21.5	1.44 (1.24, 3.11)	1.95 (1.14, 3.34)
>20	6	2.1	12	4.3	1.64 (0.89, 6.60)	2.49 (0.79, 7.82)
Unknown	2	0.8	7	2.6		
Partner's Smoking						
No	120	43.0	130	46.6	1	
Yes	129	46.2	114	40.8	0.82 (0.58, 1.16)	
Unknown	30	10.8	35	12.6		
Number of Cigarettes/Day (partner)						
0	120	43.0	130	46.6	1	
1–10	39	14.0	36	12.9	0.85 (0.50, 1.43)	
11–20	73	26.2	62	22.2	0.78 (0.52, 1.19)	
>20	17	6.0	16	5.7	0.93 (0.42, 1.80)	
Unknown	30	10.8	35	12.6		

*Adjusted on maternity hospital, maternal age, appendectomy, prior EP, prior tubal surgery, prior spontaneous abortion, previous use of intrauterine device (IUD), pelvic inflammatory disease, induced conception cycle, use of combined estroprogestative pill, IUD and progestative micropill at the time of conception.

†Smoking status at the time of conception.

§When smoking (number of cigarettes/day) was modeled as a continuous variable, the OR associated with smoking x + 1 cigarettes/day versus smoking x cigarettes/day was 1.03 (1.01 − 1.05).

link between ectopic pregnancy and smoking should therefore be considered as a public health issue. ☐

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Effects of a Worksite Nonsmoking Policy: Evidence for Increased Cessation

Glorian Sorsensen, PhD, MPH, Nancy Rigotti, MD, Amy Rosen, PhD, John Pinney, and Ray Prible, MD

ABSTRACT

A random sample of employees was surveyed 20 months after a nonsmoking policy was implemented at the New England Telephone Company in 1986. Overall, 21 percent of respondents who were smoking at the time they heard about the policy had quit smoking; 42 percent of quitters said they stopped smoking because of the policy. Cessation was highest among those who reported less smoke in their work area, but was not related to participation in cessation programs. This study suggests that worksite nonsmoking policies may have favorable effects on smoking cessation. (*Am J Public Health* 1991; 81:202–204)

Introduction

As evidence of health risks associated with passive smoking has accumulated, more worksites have adopted policies that restrict smoking.[1] Studies evaluating these policies, largely in health care facilities, have reported that nonsmoking policies are generally well-received, improve air quality, and reduce cigarette consumption.[2-6] However, there is little evidence that nonsmoking policies stimulate smoking cessation, and the generalizability of these findings to other worksites is not known.

On July 1, 1985, the New England Telephone Company, which employed 27,374 persons at approximately 600 sites, announced a company-wide policy restricting smoking. Beginning September 1, 1985, smoking was banned in conference rooms and classrooms and restricted in cafeterias and lounges. On March 1, 1986, smoking was prohibited in all work areas, including individual offices. Smoking areas were designated in cafeterias, lounges, hallways, and restrooms. A full-time field manager was appointed for 18 months to facilitate the implementation and enforcement of the policy. Free onsite smoking cessation classes were offered. This study examines the effects of this smoking policy on employee smoking behavior and perceived air quality.

Methods

Employees were surveyed in November 1987. The self-administered anonymous survey was distributed through company mail to a stratified random sample of 1,599 employees, including 892 nonmanagers (4.3 percent of technicians, operators, and secretarial/clerical staff), 494 (7.7 percent) lower level managers, and 213 (49.9 percent) upper level managers. The overall response rate was 74.5 percent (N = 1,192): 65 percent among nonmanagers, 86 percent among lower level managers, and 84 percent among upper level managers (three responses to job category were missing). Analysis is limited to data from the 1,120 (94 percent) respondents employed when the policy was implemented.

Respondents were asked about their current smoking status and smoking status when they first became aware that the company had or would be establishing a policy about smoking. Quitters were defined as those who were smokers when they first became aware of the policy and were not smoking when surveyed. Smokers included those smoking cigarettes, pipes, or cigars. As an indicator of compliance, respondents were asked about air quality in work and nonwork areas.

Relationships among these variables were examined using chi square analysis. The results are reported by job category wherever significant differences by job status are found. Multiple logistic regression was used to identify predictors of smoking status.[7] Analyses were conducted using the software program SYSTAT.[8]

Results

Characteristics of the sample are presented in Table 1. Overall, 79 (21 percent) of the 375 respondents who were smokers when they first became aware of the policy said they were not smoking by the time of the survey 20 months later, including 15 percent of nonmanagers, 25 percent of lower level managers, and 32 percent of upper level managers (chi square = 8.6, p < .01). Omitting short term quitters (N = 13), 18 percent had quit for at least three months. Differences in smoking status by job category are not explained by age or

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Address reprint requests to Glorian Sorensen, PhD, MPH, Division of Preventive and Behavioral Medicine, University of Massachusetts Medical School, 55 Lake Avenue North, Worcester, MA 01655. Drs. Rigotti and Rosen, and Mr. Pinney are with the Institute for the Study of Smoking Behavior and Policy, Kennedy School of Government, Harvard University; Dr. Prible is with New England Telephone, Boston. This paper, submitted to the Journal February 12, 1990, was revised and accepted for publication September 5, 1990.

TABLE 1—Characteristics of the Sample

Demographics	% Nonmanagers	% Lower Level Managers	% Upper Level Managers
Sex	(n = 524)	(n = 407)	(n = 177)
Female	52.5	41.5	16.9
Male	47.5	58.5	83.1
Age (years)	(n = 526)	(n = 410)	(n = 177)
18–29	8.0	4.4	1.1
30–39	36.1	30.0	12.0
40–49	31.2	37.1	39.0
50+	24.7	28.5	42.9
Smoking Status When Respondent First Heard About the Policy	(n = 528)	(n = 411)	(n = 177)
Current Smoker	33.7	34.8	29.9
Former Smoker	30.7	33.8	42.4
Never Smoker	35.6	31.4	27.7
Present Smoking Status	(n = 528)	(n = 411)	(n = 177)
Current Smoker	29.7	27.3	20.9
Former Smoker	35.0	42.1	52.4
Never Smoker	35.2	30.7	27.1

Note: A total of 79 persons who were smoking when they first heard about the policy were not smoking when the survey was administered. Also, eight who reported former smoking and three who reported never smoking when they heard about the policy reported current smoking at the time of the survey.

sex, based on logistic regression analyses. Results of a multiple logistic regression analysis indicate that smoking cessation was more common among those reporting less smoke in work and nonwork areas (Table 2).

The policy was effective in reducing reported exposure to smoke in work areas but not in nonwork areas (Table 3). These results varied significantly by smoking status but not by job category.

Employees classified as smokers were asked: "Which of the following best describes the effect the company's policy about smoking at work has had on your smoking?" Of the 375 smokers, 32 (9 percent) reported that they quit smoking due to the policy, including 20 percent of upper level managers, 9 percent of lower level managers, and 6 percent of nonmanagers. This represents 42 percent of those who quit. In addition, 113 of 375 (32 percent) said that they reduced the number of cigarettes smoked as a result of the policy, including 36 percent of upper level managers, 34 percent of lower level managers, and 29 percent of nonmanagers.

Overall, 38 percent of respondents classified as smokers received some kind of cessation-related assistance. Smokers seeking help were no more likely to quit than those not seeking help (21 percent vs 20 percent). This result did not differ significantly by job category.

TABLE 2—Logistic Regression: Smoking Cessation,[1] with Independent Variables

Variable	Odds Ratio (95% CI)
Job Status Dummy1[2] (non-managers)	1.34 (.50, 3.54)
Job Status Dummy2[2] (managers)	.83 (.32, 2.11)
Number of Cigarettes	.98 (.95, 1.00)
Smoke in Work Areas[3]	1.42 (1.04, 1.94)
Smoke in Non-Working Areas[4]	1.42 (1.03, 1.95)
Age	.82 (.58, 1.16)
Sex	1.13 (.61, 2.09)

Model chi-square 23.89, df = 7 Significant at .01 level

1. Smoking Cessation among those who were smokers when they first heard about the policy.
   1 = Current Smoker
   2 = Former Smoker
2. Upper level managers serve as the reference category for these dummy variables.
3. "How often do you notice smoky air *in your immediate work area*?"
4. "How often do you notice smoky air *in non-work areas* like hallways, rest rooms, etc?"

Discussion

A major finding of this study was the overall quit rate of 21 percent over 20 months which is markedly higher than the expected population quit rate of 2 to 5 percent per year.[9,10] In addition, 42 percent of quitters said their cessation was a direct effect of the policy. Cessation was associated with working in areas of good policy compliance, as measured by perceived air quality, but not with participating in cessation classes.

The strengths of this study include data collection from a random sample of employees and retrospective cohort data on cessation. The response rate to the survey was high (75 percent) although it was higher among managers than nonmanagers. Limitations of the study include retrospectively collected data, which risks recall bias. However, a recent study reported 87 percent agreement between recall of smoking status 20 years ago and longitudinal records.[11] Assessment of smoking cessation was based on self-report, but anonymity of responses gave little motive to distort answers. Differing employee turnover rates between smokers and nonsmokers may have inflated cessation rates; however, only 1 percent of managers surveyed knew of any employee who had left the company due to the policy, and the smoking rate of new hires was higher than in established employees. Finally, there was no comparison company to control for secular trends and our results are compared only with reports from population-wide data.

These findings differ from earlier reports from studies in health care settings, which found no relationship between smoking policies and cessation. Findings from health care settings may not apply to other settings; smokers in health care settings may be more aware of the dangers of smoking and consist of a hard core group more resistant to change. These findings suggest that a well implemented worksite smoking policy, which is fully supported by management and accompanied by cessation classes, may be followed by apparent increases in smoking cessation by employees. ☐

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TABLE 3—Effect of Smoking Policy on Air Quality

	% Current Smoker	% Former Smoker	% Never Smoker	Total
In Work Areas[1] (p = .0001)	(n = 300)	(n = 445)	(n = 350)	(n = 1095)
Improved Air Quality	55.7	71.7	76.0	68.7
Worsened Air Quality	.7	2.0	3.7	2.2
No Effect	43.7	26.3	20.3	29.1
Total	100.0	100.0	100.0	100.0
In Non-work Areas[2] (p = .0001)	(n = 297)	(n = 436)	(n = 340)	(n = 1081)
Improved Air Quality	22.2	23.4	19.0	21.7
Worsened Air Quality	32.7	50.7	55.5	47.3
No Effect	45.1	25.9	25.6	31.1
Total	100.0	100.0	100.0	100.0

1) "What effect do you think the policy about smoking at work has had on the air quality in *your immediate work area*?"

2) "What effect do you think the policy about smoking at work has had on the air quality *in non-work areas like hallways, rest rooms, etc?*"

Galvani, Director of Operations, Wellesley Dictrict; Robert B. Rand, Associate Director, Opinion Research; Donald Harney, Director of Medical Administration; Margaret Barry, Associate Director of the Employee Assistance program; and Elizabeth Warwick, Manager, Medical Liaison for Operator Services. The authors would also like to thank Drs. Lois Biener, Michael Eriksen, Jan Hitchcock, Judith Ockene, and Michael Stoto for their helpful reviews. This study was supported in part by a grant from the W. K. Kellogg Foundation.

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Are People More Health Conscious? A Longitudinal Study of One Community

Minoo Niknian, PhD, R. Craig Lefebvre, PhD, and Richard A. Carleton, MD

ABSTRACT

Secular changes in cardiovascular health awareness, knowledge and behavior were observed in four biennial cross-sectional surveys and a cohort survey in a New England community. These changes are not related to more health promotion activities in the social milieu of respondents, but are more likely due to national mass media health campaigns, the effects of which may influence outcomes of community-based cardiovascular disease prevention studies. (*Am J Public Health* 1991; 81:203–205)

Introduction

Communication campaigns designed with the objective of improving the health status of a target population have become more prevalent in the past two decades.[1-8] Many of these campaigns have been based on the assumption that people will behave in prescribed ways given sufficient information and motivation to support such decisions. An understanding of how these national campaigns are currently affecting awareness, knowledge and behavior related to cardiovascular disease (CVD) and its risk factors is important for studies that will examine risk factor changes and morbidity/mortality in study populations.

The setting for this work is the Pawtucket Heart Health Program.[5,9] In this paper, the question of whether residents of the comparison community have become more health conscious—especially with respect to CVD—is the focus.

Methods

The Pawtucket Health Heart Program conducts a series of biennial surveys to assess relevant behaviors and cardiovascular risk factors in the study communities. Each of these surveys includes interviews and physiologic assessments of approximately 1,250 citizens between the ages of 18 and 65 in each of two New England cities, and spans the entire year in order to control for seasonal differences in physiologic risk factors. Households (defined as in the United States Census) are randomly sampled in each city using current listings known to be at least 95 percent complete. A single respondent from age-eligible members of the household is selected using tables adapted from Kish[10] and Deming.[11] A 15–20 minute interview includes questions about diet, exercise, smoking, CVD knowledge, availability of CVD-related programs, and sociodemographic characteristics. Measurements of height and weight, pulse, and blood pressure, and a 30 ml blood sample for later analysis of total cholesterol, triglycerides and HDL-cholesterol, are also obtained. The survey methods and their reliabilities are discussed in detail elsewhere.[12]

The data presented in this paper were collected in the four cross-sectional surveys conducted in the comparison city in 1981–82, 1983–84, 1985–86, and 1987–88. In addition, respondents to survey 1 were resurveyed after a lapse of five years: 61.3 percent completed the second examination. Each respondent was reexamined within one calendar month of the date of their first examination.

The percentage of participants answering favorably to questions were considered as dependent variables for surveys. Physiologic variables will not be analyzed until the final surveys are completed. Potential covariates employed in the analyses were: age, gender, education, language spoken at home, ethnicity, marital status, employment status, and household income level.

The objectives of this report are to document awareness, knowledge and behavior changes over time and determine trends. To measure the linear trend, the slope was calculated by weighted linear regression for each dependent variable, and tested with the Chi-square statistic.[13,14]

To adjust for potential covariate effects, eight groups were constructed. The cutpoints were investigated carefully so that: (a) each group had enough observations, and (b) the differences between distribution of surveys were preserved. The linear trend was calculated for each group and then an adjusted overall measure of linear trend across all groups was taken as the weighted average of the eight trends.[15]

From the Department of Community Health, Brown University Program in Medicine (Niknian, Lefebvre); and the Department of Medicine, Memorial Hospital of Rhode Island, and Brown University Program in Medicine (Carleton).

Address reprint requests to Minoo Niknian, PhD, Biometrics Department, Warner-Lambert Company, 2800 Plymouth Road, P.O. Box 1047, Ann Arbor, MI 48106-1047. This paper, submitted to the Journal January 26, 1990, was revised and accepted for publication October 10, 1990.

TABLE 1—Results of Cross-sectional Survey Questions Regarding Awareness and Behavior Related to CVD Prevention, and Results of Linear Trend

Variables	S1 (1280)	S2 (1442)	S3 (1484)	S4 (864)[a]	X^2	b	SE (b)
Is it possible to prevent CVD	83.8	87.4	88.5	91.9	31.816**	.0250	4.432×10^{-3}
Methods(s) Mentioned:							
▶ Quit Smoking	44.4	41.3	44.9	45.1	.688	.0059	7.113×10^{-3}
▶ Cut salt intake	13.0	15.1	16.6	18.5	11.963**	.0181	5.233×10^{-3}
▶ Cut fat intake	27.8	45.1	55.3	63.5	272.970**	.1189	7.197×10^{-3}
▶ Cut calories	8.5	4.8	6.1	6.4	2.288	−.0053	3.504×10^{-3}
▶ Increase exercise	39.7	45.0	53.0	52.8	47.890**	.0498	7.196×10^{-3}
▶ Control weight	32.0	32.0	28.6	27.6	6.643*	−.0170	6.596×10^{-3}
▶ Control blood pressure	10.8	10.5	11.2	8.4	1.439	−.0053	4.418×10^{-3}
▶ Control stress	29.6	29.7	30.1	24.7	3.416	−.0121	6.547×10^{-3}
Do not know own risk of CVD	8.9	9.1	8.8	8.2	.292	−.0021	3.886×10^{-3}
Do not know own BP level	8.4	9.9	10.8	11.0	5.023*	.0091	4.051×10^{-3}
Do not know own CHOL level	35.0	37.7	35.9	25.3	16.194**	−.0258	6.411×10^{-3}
In last year, have:							
▶ Had CHOL measured	11.6	12.3	13.2	16.3	9.597**	.0141	4.551×10^{-3}
▶ Attempted weight loss	48.8	50.7	51.1	46.4	.353	−.0040	6.732×10^{-3}
▶ Maintained loss	56.6	54.9	59.9	57.1	.890	.0091	9.646×10^{-3}
▶ Increased exercise	30.0	32.8	32.7	34.5	4.402*	.0133	6.315×10^{-3}
▶ Maintained increase	59.1	57.7	55.8	59.4	.063	−0.0030	11.952×10^{-3}
Current cigarette smoker	42.8	39.9	39.4	36.0	9.371**	−.0202	6.599×10^{-3}
▶ Attempt to quit in past 12 mos.	41.1	45.6	48.5	44.4	2.704	.0175	10.642×10^{-3}
Live with smoker	38.4	38.8	36.5	35.3	3.079	−.0115	6.554×10^{-3}
Limit salt intake	65.8	71.2	67.9	67.6	.187	.0027	6.244×10^{-3}
Limit fat intake	63.3	66.5	68.5	66.9	5.288*	.0147	6.379×10^{-3}
Never encourage others to							
▶ Stop smoking	24.1	21.3	21.3	21.9	1.745	−.0074	5.602×10^{-3}
▶ Lose weight	28.6	24.2	27.0	26.2	.556	−.0044	5.901×10^{-3}
▶ Cut salt/fat intake	38.0	28.4	31.5	29.9	12.293**	−.0221	6.297×10^{-3}
▶ Exercise more	36.1	30.7	32.4	29.5	7.673**	−.0175	6.314×10^{-3}
Membership organization offered health programs in past year	16.5	17.1	17.7	18.9	1.986	.0076	5.393×10^{-3}
▶ Nutrition	19.7	17.3	17.5	18.9	.041	−.0027	13.334×10^{-3}
▶ Stress	4.8	9.7	18.8	12.1	10.770**	.0358	10.909×10^{-3}
▶ Weight	25.0	18.9	31.2	27.7	2.622	.0240	14.822×10^{-3}
▶ Exercise	35.1	40.7	38.0	41.2	.780	.0146	16.531×10^{-3}
▶ Blood Pressure	17.0	18.1	13.7	6.8	8.076**	−.0339	11.933×10^{-3}
▶ Smoking	16.0	15.9	26.5	31.8	17.376**	.0586	14.058×10^{-3}

[a] Due to budgetary constraints, S4 was extended to two years. Only data from the first half of the survey (1987–88) are reported.
*p ≤ .05
**p ≤ .01
Results expressed as percentages unless otherwise noted.

Results

A total of 5,070 persons participated in one of the cross-sectional surveys in the comparison city. Preliminary analyses were performed for each covariate. Gender distribution and marital status were similar across the four surveys with females comprising 58 percent of the overall sample; approximately 60 percent of respondents were married.

Significant differences among the four surveys were noted for age group distribution, years of education and proportion of employed. Survey participants at the last survey tended to be younger, better educated and more often employed compared to the first one.

Persons of Portuguese descent comprised about 45 percent of the sample in each survey, and only 73 percent of the sampled population spoke English. There was no indication of significant changes in these variables across the surveys.

Using the weighted regression analyses to detect linear trends in CVD knowledge and behavior, a number of significant effects were found (see Table 1). When each of these linear relationships was adjusted for the covariates (i.e. age, education level and employment status), similar trends were found for each variable.*

The expected changes in the population from regression analysis of changes in the cross-sectional surveys were compared with the actual changes found among cohort survey respondents.* The data indicate that, in most cases, the observed change in the cohort exceeded the expected value derived from the cross-sectional surveys.

Discussion

These results support the conclusion that important secular changes are occurring in cardiovascular health awareness, knowledge and behavior. It appears that such effects are not due to the increased presence of prevention programs either at participants' worksites or religious and social organizations and are likely attributable to programs mediated through electronic and print mass media.

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*Data available on request to authors.

Respondents in the cohort survey demonstrated greater changes in health promotion knowledge and behavior than expected from cross-sectional survey data. This may reflect a self-selection bias among those persons who were reexamined, although the first examination may also have induced more health consciousness among this group.

The large proportion of persons of Portuguese heritage is a factor that limits generalizing these data to other sections of the region or country. The high proportion of respondents who spoke a foreign language at home raises the possibility that our estimates of the "secular trend" for CVD prevention-related awareness, knowledge and behavior are conservative. In more homogeneous, affluent, and educated communities, even greater effects may have occurred in the six-year period studied here.[16]

These observations support other studies of changes in awareness, knowledge and behavior among the general public.[8] That these changes are occurring in a city that serves as a comparison area for a population-based CVD research and demonstration project may reduce the power to detect the effects of intervention in the intervention community.[17-21] □

Acknowledgments

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Sexual Risk Behaviors, AIDS Knowledge, and Beliefs about AIDS among Runaways

Mary Jane Rotheram-Borus, PhD, and Cheryl Koopman, PhD

ABSTRACT

Sexual risk behaviors, knowledge of acquired immunodeficiency syndrome (AIDS), and beliefs about AIDS prevention were examined among 126 runaways. In the previous 3 months, 65 percent of youths had been sexually active. Among the sexually active runaways, males reported a median of 2.7 partners and females reported 1.3 partners, and only 18 percent reported consistent condom use. Runaways demonstrated moderately high AIDS knowledge and beliefs endorsing AIDS prevention. Condom use and abstinence were directly related to beliefs about preventing AIDS. (*Am J Public Health* 1991;81:206–208)

Introduction

In 1988, runaways in New York City (NYC) were found to have human immunodeficiency virus type 1 (HIV-1) seropositivity rate of 6.7 percent.[1] This finding identified runaway youths as a high-risk group for HIV-1 infection. To design effective prevention programs, more information is needed on the sexual risk behaviors of these youths. However, empirical data documenting their risk behaviors are limited to a few items in epidemiological studies of runaways.[2–4]

Knowledge of AIDS and beliefs about AIDS prevention are relevant to reducing HIV-1 infection.[5,6] Therefore, the goals of this paper were to document runaways' current sexual risk behaviors, knowledge, and beliefs about preventing AIDS, and to examine the relationships among these findings.

Methods

Sample

The sample includes consecutive admissions to two of three residential shelters for runaway youths in Manhattan and Brooklyn over an 8-month period in 1988–89, for a total of 64 males and 66 females. They were recruited by two young adult Black interviewers, one male and one female, of whom both had intensive training in psychosexual assessment.[7] Two runaway males refused to participate in the study. Each sample contained primarily Black (65 percent) and Hispanic youths (28 percent). They ranged in age from 11 to 17, and 70 percent were born in NYC. The average age was 15.7 (SD = 1.5), and the median education completed was sophomore year of high school. Only 22 percent of runaways had ever lived "on the street"; however, 75 percent had not lived with a parent or guardian for at least 2 weeks.

Sexual Risk Behavior Assessment

A semistructured interview guide, *the Sexual Risk Behavior Assessment Schedule–Youth* (SERBAS),[8] was used to elicit information about sexual behavior. In this study, assessments of current sexual activity were limited to acts in which the exchange of semen or vaginal fluid was possible (i.e., oral, anal, and penile-vaginal intercourse), although these acts may not be of equal risk. Sexual behaviors assessed over the previous 3 months included abstinence, number of sexual partners, number of sexual encounters, and frequency of condom use (assessed on a 4-point scale based on number of sexual encounters divided by instances of condom use). Test-retest reliability conducted on items readministered to 22 runaways after 1 week was high (r ranged from 0.71 to 1.00).

General Knowledge about AIDS

A true-false instrument containing 62 items was used to assess general AIDS knowledge. The process of developing the instrument helped to address its validity, resulting in high internal consistency and test-retest reliability, as described elsewhere.[9] This instrument covers six domains: definitions, outcomes, risk behavior, transmission, prevention, and HIV-1 testing.

Beliefs about Preventing AIDS

Beliefs about AIDS and safer sex practices were assessed with 39 Likert-type items. This measure was developed in parallel procedures with the AIDS knowledge instrument. Previous research[9] using this instrument with runaway adolescents showed that it has high internal consistency. It covers five domains: peer support for safe acts, self-efficacy, self-control, perceived threat, and expectation of acting to prevent pregnancy, which is included because some of

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Address reprint requests to Mary Jane Rotheram-Borus, Box 29, Division of Child Psychiatry, Columbia University, 722 W. 168th Street, New York, NY 10032. Drs. Rotheram-Borus and Koopman are affiliated with the HIV Center for Clinical and Behavioral Studies, New York State Psychiatric Institute, Columbia-Presbyterian Medical Center, Columbia University. This paper, submitted to the Journal February 1, 1990, was revised and accepted for publication June 13, 1990.

TABLE 1—Frequencies of High-Risk Sexual Behavior during the Past 3 Months among Sexually Active Runaways[a]

Behavior	Males (N = 44)	Females (N = 37)
Sexually active	70%	59%
Number of partners		
Median	2.7	1.3
Mean ± SD	4.7 ± 6.0	2.0 ± 3.1
Number of sexual occasions		
Median	9.0	5.3
Mean ± SD	14.9 ± 20.6	9.9 ± 12.8
Condom use, percent		
Never	27.0	47.2
Sometimes	35.1	13.9
Most of the time	21.6	19.4
Always	16.2	19.4

[a] Sexual activity includes oral, anal, and vaginal intercourse. Data apply only to sexually active runaways. Thirty percent of the 63 male runaways and 41 percent of the female runaways in the total sample reported no sexual activity.

TABLE 2—Means and Standard Deviations of General Knowledge of AIDS and Positive Beliefs about Preventing AIDS among Male and Female Runaways

Variable	Males (N = 61)	Females (N = 64)
General knowledge	73% ± 12.9%	74% ± 11.6%
Belief domains		
Total positive beliefs	120.2 ± 13.6	122.3 ± 14.5
Peer support	16.1 ± 3.2	16.0 ± 3.8
Self-efficacy	63.4 ± 8.4	63.3 ± 7.7
Perceived threat	10.4 ± 2.0	10.4 ± 1.8
Self-control	16.2 ± 2.8	17.0 ± 1.8
Prevent pregnancy	14.7 ± 3.0	15.6 ± 3.8

the major strategies for preventing pregnancy (abstinence, avoiding sexual intercourse, and condom use) also reduce risk for HIV-1 infection.

Data Analysis

The frequencies of high-risk sexual behaviors were calculated for male and female runaways. One male and three females who reported exclusively homosexual activity were removed from analyses of sexual risk behaviors because their sexual behavior was not comparable to that of the heterosexual runaways. Chi square tests were computed on sexual risk behaviors, splitting the data on the number of sexual partners and number of sexual encounters at the joint median. T-tests examined gender differences on AIDS knowledge and on overall and domain scores for beliefs about AIDS prevention. Regressions were computed between overall knowledge and belief domain scores with sexual risk behaviors.

Results

The majority of these high-risk youths had been sexually active in the previous 3 months (see Table 1). Sexually active males reported a median of 2.7 sexual partners and 9.0 sexual encounters, and sexually active females reported a median of 1.3 sexual partners and 5.3 sexual encounters; 65 percent of the females and 32 percent of the males had only one partner. Among the sexually active youths, consistent condom use was low, although 47 percent of males and 44 percent of females reported using a condom the last time they had sexual intercourse. No runaways reported bisexual activity in the previous 3 months.

Table 2 shows the overall AIDS knowledge scores and overall and domain-specific beliefs of male runaways and female runaways. Youths showed moderately high general knowledge ($\bar{X}$ = 45.4, SD = 7.6, [a score that reflected 73 percent accurate knowledge]). Positive beliefs about preventing AIDS were also high ($\bar{X}$ = 121.4, SD = 14.1, [reflecting a 70 percent positive endorsement of safe acts]). There were no significant differences between the male and the female runaways on these measures.

Although the sexual risk behaviors were not significantly related to overall AIDS knowledge, condom use showed a weak inverse relation to perceived threat (B = −.17, 95 percent CI = −.33, −.01), and expecting to act to prevent pregnancy was positively associated with condom use (B = .14, 95 percent CI = .07, .22) and with abstinence (B = .05, 95 percent CI = .02, .07).

Discussion

Because previous research[3] has shown that intravenous (IV) drug use is not high among runaways in NYC, sexual activity is likely to be their primary risk behavior for HIV-1 transmission. The runaways' sexual risk behaviors, AIDS knowledge, and beliefs about AIDS do not provide a sufficient explanation for the high seroprevalence rate reported in 1988 (6.7 percent). Even though the runaways evaluated in this study were drawn from a residential shelter other than the one where the 6.7 percent seroprevalence rate was found, there is no reason to believe that the youths in our sample differ from those in the prevalence study. They were similar in age, ethnicity, and socioeconomic status, and many had been seen at the shelter where the seroprevalence study was conducted.

While most (65 percent) had been sexually active during the previous 3 months, the rate was not dramatically different from the percentage of a comparable sample of high school students who had been sexually active in the previous 6 months (58 percent).[10] Although condom use was relatively inconsistent, again, it was not different from condom use among adolescents in general.[11] Sexually active runaway girls typically have reported only one partner in the past 3 months. Furthermore, runaways' moderately high knowledge and positive beliefs about AIDS prevention were similar to those of high school samples.[10,12] This pattern of results suggests the need for more detailed studies examining the social networks from which these youths draw their sexual partners. Future research must also examine this population's drug use, including IV drug use, needle sharing, and crack use, as well as their exchange of sex for drugs or money.

These data do point to the importance of developing specialized programs for male runaways. Males have significantly more sexual partners compared with runaway females, consistent with findings on the number of lifetime sexual partners in a national sample of male adolescents.[13] However, the percentage of males in the study who used a condom the last time they had sexual intercourse (47 percent), was less than reported by adolescent males in a national survey (58 percent).[14]

Consistent with previous research,[15] we found moderate relationships between beliefs about preventing AIDS and sexual risk behaviors. Although as a group, the runaways had a fairly high knowledge of AIDS and had positive beliefs about AIDS, some youths were relatively uninformed. Interventions with these youths may impact behavior. ☐

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Security Measures for AIDS and HIV

Carl G. Torres, PhD, Mark E. Turner, MPH, John R. Harkess, MD, and Gregory R. Istre, MD

ABSTRACT

This study describes the measures being taken by AIDS surveillance offices across the country to ensure the security of information regarding patients with AIDS and HIV infection. Security measures were evaluated according to the cumulative number of AIDS cases reported, whether partner notification services were provided, and whether HIV seropositive reporting by name was also required. This study showed that public health departments have taken extra steps to ensure the security of AIDS and HIV data. (*Am J Public Health* 1991; 81:208–209)

Introduction

Public health surveillance for acquired immunodeficiency syndrome (AIDS) and human immunodeficiency virus (HIV) infection has accentuated the need for confidentiality and security of health records. This study describes the measures being taken across the country to ensure the security of information regarding patients with AIDS and HIV infection reported to public health departments. Security measures were evaluated according to the cumulative number of AIDS cases reported, whether partner notification services were provided, and whether HIV seropositive reporting by name was also required.

Methods

The study population comprised all 66 AIDS surveillance offices responsible for reporting to the Centers for Disease Control as of July 1, 1989. These offices were located in 13 city/county health departments, all 50 state health departments, the District of Columbia, Puerto Rico, and Guam. Confidential questionnaires solic-

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From the Epidemiology Service, Oklahoma State Department of Health. Address reprint requests to Mark E. Turner, MPH, Epidemiology Service, Oklahoma State Department of Health, P.O. Box 53551, 1000 N.E. Tenth Street, Oklahoma City, OK 73152.

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TABLE 1—Security Measures for Case Reports and Computerized Data with Personal Identifiers in 63 AIDS Surveillance Offices, 1989

	%	N	%
Case Report Security Measures			
Reports stored in a locked file cabinet or safe	100		
Additional locking system	23		
Fire-proof cabinet	44		
File cabinet in a restricted area	87		
Offices with any one or more of above three			91
Alarm system for the storage area	8		
Alarm system for the building	54		
Security guard in the building during non-working hours	70		
Offices with any one or more of above three			83
Median number of employees with access to restricted area	7	(2–30)*	
Median number of employees with access to data	3	(1–15)*	
	%	**Ñ**	**%**
Computerized Data Security Measures for Offices with Personal Identifiers in Their Computer System, N = 54 (86%)			
Use security key lock	74		
Computer in a restricted area	78		
Computer bolted in a permanent location	17		
Use a removable hard drive system	15		
Use special software to control access to the computer	54		
Computer data encrypted	17		
Offices with any one or more of above six			98
Alarm system for computer area	7		
Alarm system for the building where computer is located	50		
Offices with any one or more of above two			52
Median number of employees with access to restricted area	6	(2–30)*	
Median number of employees with access to computer data	3	(1–5)*	
Median number of identification codes and passwords used	2	(1–3)*	

*Range

iting information regarding security measures used to protect AIDS and HIV data were sent in mid-July 1989. A reminder letter was sent in mid-August 1989 and reminder telephone calls were placed during the last week of August 1989. Security measures were examined by the number of cumulative AIDS cases reported to the specific office as of June 30, 1989 (less than 500, between 500 and 1,500, and greater than 1,500), whether provider referral partner notification (PRPN) was offered, and whether HIV seropositive reporting by name was required.

Results

Questionnaires were returned by 63 (95 percent) of 66 AIDS surveillance offices. All offices maintained their records in locked cabinets. In addition, an array of other measures were employed to protect written and computer records (Table 1). Where all HIV infections were reportable, 95 percent of the offices handled HIV reports in the same way as their AIDS case data. There were 26 offices with fewer than 500 AIDS cases reported, 20 offices with 500 to 1,500 cases, and 17 offices with over 1,500 cases. Forty-six offices (73 percent) provided PRPN. The median number of AIDS cases reported in offices with PRPN was 499, while the number for offices without this service was 1,400. Reporting of all HIV infections was required by 21 offices. The median number of AIDS cases in the 21 offices was 550, while in the offices without HIV reporting the median number was 1,273. The proportion of security measures utilized did not differ significantly with respect to the cumulative number of AIDS cases reported, whether PRPN was offered, or whether all HIV infections were reportable.

Discussion

This study showed that public health departments have taken steps to ensure the physical security of records regarding patients with AIDS and HIV infection. As a minimum, records were maintained in locked cabinets, however, most offices had instituted additional safeguards, such as restricting access to areas where AIDS records are stored, bar locks, alarms, and computer security systems.

No statistically significant differences were found in the proportion of security measures with respect to the cumulative number of AIDS cases reported, whether PRPN was offered or whether reporting of HIV infection was required. This indicates that the implementation of PRPN and HIV reporting is being done without compromising the physical security of records. The results of this study provide a tool with which public health departments and others holding highly sensitive personal health data can appraise the measures taken to ensure the security of their records. Although this study does not address procedures being taken to ensure security of data among field staff or in settings outside public health departments, it does indicate that public health officials are taking extra measures to ensure the physical security of data regarding patients with AIDS and HIV infection. ☐

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Discharges against Medical Advice at Regional Acute Care Hospitals

David Barton Smith, PhD, and Joel Leon Telles, PhD

ABSTRACT

Data from 67 acute care hospitals in the Philadelphia metropolitan area indicate that there were 7,613 discharges against medical advice (AMA) in fiscal year 1987, or 1.20 percent of all discharges that year. Diagnosis-related group (DRG), type of insurance, and sex had independent effects on the rate of AMA discharges. Urban community hospitals had the highest percent of AMA discharges. Previous studies, done in teaching facilities, may have underestimated the rate of AMA discharges. (*Am J Public Health* 1991; 81:210–213)

Introduction

Discharges against medical advice (AMA) represent a frustrating problem for the individual physician and hospital, and an overlooked indicator of potential problems facing the organization of public health services in a region. Little systematic analysis of the subject has been published. Most reports have been anecdotal or limited to specific diagnostic groups within a few selected institutions.[1-3] The only published general estimate, based on a 2-year review of records in a major teaching institution, reported that patients who were discharged AMA accounted for about 0.8 percent of discharges.[4] We analyzed regional data to describe regional patterns and, because larger numbers were involved, to help clarify the relative influence of factors associated with variation in AMA rates.

Methods

A systematic description of patients discharged against medical advice from acute care hospitals was generated from the comprehensive regional inpatient data base constructed directly from hospital abstract tapes supplied to the Delaware Valley Hospital Council. We analyzed the data on the 635,897 patients discharged from all 67 acute care hospitals in the Greater Philadelphia area between July 1, 1986, and June 30, 1987. We used a linear modeling approach, applying the categorical modeling procedure in the SAS statistical package, to estimate the independent effects of diagnosis-related group (DRG), payment method, and sex on the percent of AMA discharges.[5,6]

Results

A total of 7,613 cases, or 1.20 percent of inpatient discharges, were listed as discharged AMA. Males were more than twice as likely as females to be discharged AMA (Table 1). Patients who were covered by the Medicaid program or who had no insurance were three times as likely to be discharged in this manner.

AMA discharges were concentrated in a few diagnostic categories. Table 2 lists the 20 DRGs with the highest frequency of AMA discharges according to the Medicare prospective-payment grouper procedure in effect October 1, 1987, to September 30, 1988.[7] These 20 DRGs accounted for 60.7 percent of AMA discharges. Major diagnostic category (MDC) 20 (DRG 433-437), alcohol and drug use and alcohol- and drug-induced organic mental disorders accounted for 26.6 percent of AMA discharges and was the MDC with highest percentage of patients with such discharges (14.6 percent).

The average length of stay for AMA discharges was about half the average length of stay for all discharges in most of the 20 DRGs. The DRG case weight for all AMA discharges was 0.77, compared with 0.99 for all discharges in the region, suggesting that they were somewhat less severe.[8] However, it is also possible that many such discharges led to later readmissions, as noted in studies of psychiatric facilities.[9]

Among the 20 most frequent AMA discharge diagnoses, diagnosis, insurance status, and sex of the patient had statistically significant ($p < .001$) independent main effects. The Medicaid and self-pay discharges represented underinsured groups in the Philadelphia area with similar patterns of use, and they have been combined in this analysis. The estimate of the percent of AMA discharges for any combination of these three variables can be calculated by adding the individual effects.

Thus, in Table 3, the estimate for the percent of AMA discharges for male Medicaid cases with DRG 430 (psychoses) is 8.00 percent (INTERCEPT [4.12 percent] + DRG [2.19 percent] + MEDICAID/SELF-PAY [1.35 percent] + MALE [.35 percent]. The actual percent of AMA discharges for this group of cases was 7.89 percent, and the standard error for the es-

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Address reprint requests to David B. Smith, Department of Health Administration (006-00), Temple University, Philadelphia PA 19122. Dr. Smith is with the Department of Health Administration, Temple University and Dr. Telles is with the Delaware Valley Hospital Council, Inc. This paper, submitted to the Journal January 25, 1990, was revised and accepted for publication June 12, 1990.

TABLE 1—Total Discharges against Medical Advice and Rate per 100 Discharges in the Philadelphia Metropolitan Area, by Source of Payment and Sex

	Female		Male		Total	
Payer	N	Rate/100(%)	N	Rate/100(%)	N	Rate/100(%)
Other[a]	1,281	0.45	2,039	0.91	3,320	0.65
Medicaid	1,290	2.19	2,199	5.31	3,489	3.48
Self-pay	240	1.68	564	4.51	804	3.00
Total	2,811	0.78	4,802	1.73	7,613	1.20

[a] Medicare and private insurance (Blue Cross, health maintenance organizations, and commercial insurers.

timate was 0.24 percent. For all 20 of these discharge diagnoses, estimates of the effect of the DRG category on the percent of AMA discharges ranged from 9.56 percent for substance abuse admissions falling into MDC 20 to −2.29 percent for DRG 243 (medical back problems). The effect of Medicaid or self-pay status (1.34 percent) was almost four times greater than the effect of being male (.35 percent).

AMA discharge rates also varied significantly by the geographic location and the type of facility. Suburban and teaching hospitals had the lowest rates of AMA discharges. Urban community hospitals had the highest rates. Table 4 presents the percent of AMA discharges by type of hospital for the 24 MDCs. Except for the group of specialty facilities (800 beds out of a regional total of 17,555 beds, the majority belonging to two children's hospitals), suburban facilities had the lowest AMA rates. In contrast, the urban community hospitals had the highest percent of AMA discharges (3.49 percent compared with 1.20 percent for the region as a whole). Teaching facilities tended to have lower rates of AMA discharges than community hospitals. The differences in rates of AMA discharges between urban and suburban hospitals and between community and teaching hospitals persist in almost all MDCs.

Discussion

While the analysis confirms the patterns found in the previous case studies of self-selected institutions, it places the findings into a broader perspective.[1-4] Because previous inquiries studied major teaching institutions, they may have underestimated the magnitude of the problem. The 1.2 percent of Philadelphia-area discharges that were AMA is substantially higher than the 0.8 percent found in the previously cited Peter Bent Brigham Hospital study. Urban community hospitals located in areas with the most serious problems of social disorganization (e.g., crime, drug addiction, unemployment) had the highest rates.

AMA rates do not appear to be simply a function of the disease and the nature of its treatment. Males, Medicaid patients, and patients identified as self-paying were more likely to be discharged AMA, independent of diagnosis. They represent a crude measure of therapeutic failure or, at least, failure in the provider-patient relationship. Efforts to reduce these discharges must be balanced against the individual rights of patients to refuse treatment, the legitimacy of some of the personal reasons that motivate them to leave, and the potential harm to the patient of premature discharge.

It is useful to view AMA discharges in the context of overall patterns of use and access to health care. Indeed, there is a striking symmetry between the factors

TABLE 2—AMA Rates and Length of Stay for 20 DRGs with Highest Number of AMA Discharges

DRG #	Code	N	Rate/100 (%)	Average AMA Length of Stay (LOS)	Ratio of LOS to Total[a]
MDC20	Alcohol and drug use and alcohol- and drug-induced mental disorders	2,025	14.59	3.37	0.59
430	Psychoses	686	6.13	7.79	0.48
450	Poisoning and toxic effects of drugs, age >17 without CC	250	10.35	1.20	0.36
140	Angina pectoris	196	1.38	2.60	0.42
449	Poisoning and toxic effects of drugs, age >17 with CC	154	6.58	2.37	0.26
426	Depressive neuroses	152	9.92	5.03	0.33
143	Chest pain	134	2.54	1.88	0.44
127	Heart failure and shock	101	0.72	5.14	0.49
24	Seizure and headache, age >17 with CC	98	3.62	3.20	0.29
204	Disorders of the pancreas except malignancy	95	4.50	4.58	0.33
97	Bronchitis and asthma, age >17 without CC	90	2.26	2.55	0.43
277	Cellulitis, age >17 with CC	87	3.60	4.22	0.38
294	Diabetes, age >35	78	1.89	5.71	0.55
395	Red blood cell disorders, age >17	77	2.29	4.83	0.54
243	Medical back problems	74	0.57	3.87	0.40
182	Esophagitis, gastroenteritis, and miscellaneous digestive disorders, age >17 with CC	73	1.08	3.23	0.33
174	Gastrointestinal hemorrhage, age >70	66	1.99	4.45	0.30
122	Circulatory disorders with acute myocardial infarction	64	1.13	3.95	0.38
96	Bronchitis and asthma age, >70 and/or CC	63	1.35	3.93	0.36
88	Chronic obstructive pulmonary disease	59	1.24	4.98	0.40
Top 20 AMA DRGs		4,622	3.79	3.76	0.39
All regional discharges		7,613	1.20	4.40	0.61

[a] AMA average length of stay divided by average length of stay for all discharges.

TABLE 3—Linear Model Estimates of Main Effects on Percent AMA Discharges

		Effect (%)	Standard Error (%)	AMA Estimate[a] (%)
Intercept		4.12%	0.08%	
MDC20_	Substance abuse and substance-induced organic mental disorders	9.56	0.29	13.68
430	Psychoses	2.19	0.22	6.31
450	Poisoning and toxic effects of drugs, age >17 without CC	5.85	0.57	9.97
140	Angina pectoris	−1.86	0.09	2.26
449	Poisoning and toxic effects of drugs, age >17 with CC	1.78	0.43	5.90
426	Depressive neuroses	5.88	0.71	1.00
143	Chest pain	−1.09	0.18	3.03
127	Heart failure and shock	−2.15	0.09	1.97
24	Seizure and headache, age >17 with CC	−1.20	0.24	2.92
204	Disorders of the pancreas except malignancy	−1.05	0.28	3.07
97	Bronchitis and asthma, age >17 without CC	−1.39	0.20	2.73
277	Cellulitis, age >17 with CC	−1.54	0.21	2.58
294	Diabetes, age >35	−1.64	0.17	2.48
243	Medical back problems	−2.29	0.88	1.83
395	Red blood cell disorders, age >17	−1.45	0.23	2.67
182	Esophagitis, gastroenteritis, and miscellaneous digestive disorders, age >17 with CC	−1.94	0.12	2.18
174	Gastrointestinal hemorrhage with CC	−1.69	0.17	2.43
122	Circulatory disorders with acute myocardial infarction without cardiovascular complications discharged alive	−2.05	0.13	2.07
96	Bronchitis and asthma, age >17 with CC	−2.02	0.13	2.10
Insurance (Medicaid or self-pay)		1.34	0.07	5.46
Sex (male)		0.35	0.03	4.47

[a] Intercept + Main Effect, all effects significant, p <.001

TABLE 4—Percentage of Patients Discharged AMA, by Hospital Type and Major Diagnostic Category

		Teaching Hospitals			Community					
		Suburban	City		Suburban	City				
	Major Diagnostic Category	Major	Minor	Major	Minor		Medical School	Other Specialty	Average	
1	Nervous system	0.5	1.1	1.3	1.6	1.0	2.1	1.1	0.0	1.1
2	Eye	0.0	0.0	0.7	0.6	0.1	0.7	0.2	0.0	0.2
3	Ear, nose, and throat	0.0	0.5	0.3	0.4	0.1	0.8	0.4	0.0	0.2
4	Respiratory system	0.6	1.0	1.1	1.3	0.5	2.6	1.2	0.1	1.0
5	Circulatory system	0.6	0.7	1.2	1.0	0.7	2.0	0.8	0.1	1.0
6	Digestive system	0.4	0.6	0.9	0.9	0.5	1.7	0.7	0.0	0.7
7	Hepatobiliary system and pancreas	0.7	0.7	1.8	0.7	1.0	3.7	1.3	0.0	1.4
8	Musculoskeletal system	0.2	0.2	0.6	0.7	0.3	1.1	0.4	0.1	0.5
9	Skin, subcutaneous tissue	0.5	1.1	2.2	1.5	0.5	3.5	1.5	0.0	1.4
10	Endocrine, nutritional and met.	0.5	0.8	1.5	1.1	0.8	2.5	0.8	0.1	1.1
11	Kidney and urinary tract	0.3	0.6	0.8	0.8	0.6	1.2	1.3	0.0	0.8
12	Male reproductive system	0.3	0.1	0.4	0.7	0.3	0.7	0.5	0.4	0.4
13	Female reproductive system	0.0	0.2	0.6	0.3	0.1	0.9	0.4	0.0	0.4
14	Pregnancy, childbirth	0.1	0.0	0.4	0.1	0.2	0.3	0.5	0.0	0.3
15	Newborns and neonates	0.1	0.0	0.0	0.1	0.0	0.0	0.1	0.1	0.1
16	Blood, blood forming and imun.	0.3	2.2	1.6	2.2	1.3	5.2	2.2	0.0	1.7
17	Myeloproliferative, malignancy	0.1	0.0	0.4	0.9	0.1	0.5	0.3	0.1	0.2
18	Infections and parasitic	0.5	0.0	1.5	1.7	0.7	2.8	1.3	0.3	1.1
19	Mental	4.3	4.2	5.0	3.7	7.9	5.3	7.9	1.6	6.4
20	Alcohol/drug use	7.1	14.4	13.4	7.4	12.3	18.7	7.1	0.0	14.6
21	Injuries, poisonings and toxic	2.7	4.9	4.1	6.0	2.6	7.9	1.6	0.2	3.5
22	Burns	0.4	3.6	1.7	1.0	1.0	2.9	3.3	0.0	1.2
23	Factors influencing services	0.9	1.0	0.6	1.4	0.6	3.6	0.3	0.4	0.7
24	Other	0.0	0.0	2.6	0.0	2.0	2.5	0.5	0.0	0.8
	Average	0.55	0.99	1.37	1.00	0.94	3.49	1.0	0.07	1.20

affecting "entrance" to health services (i.e., barriers and delays in seeking care and overall patterns of underutilization) and of the factors affecting AMA discharge rates, which may be viewed as inappropriate "exit" from such services.[10-14]

Since regional data systems are now widely available, it is possible to monitor these processes more systematically. These rates can be linked more directly to the relationships between medical and social problems in regional populations and to the pressures imposed on medical management by different forms of reimbursement as they change the character of inpatient care. AMA discharges represent many of the problems (e.g., those related to drugs, alcohol, psychiatric conditions, and indigent care) that public health agencies address. The percent of AMA discharges may in time take its place among other indicators of incidence rates and use of health services in a population and help document progress (or the lack of it) in solving larger social problems. ☐

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An Evaluation of the Use of the Statewide Planning and Research Cooperative System of New York State as a Resource Planning Tool for HIV Infection

Gerald I. Kaufman, PhD, MCS, George T. DiFerdinando, Jr, MD, MPH, and Sarah E. Gottesman, MS

ABSTRACT

Hospital discharge records of patients with *Pneumocystis carinii* pneumonia (PCP) in New York State were studied to determine whether cases of human immunodeficiency virus (HIV) infection were identified. We estimate that as many as 13 percent of hospitalizations of patients with PCP in 1987 and 10 percent of those in 1988 were not appropriately identified as HIV related. Identification of PCP as HIV related was a function of a hospital's volume of PCP admissions. (*Am J Public Health* 1991;81:213–215)

Introduction

One means of estimating acute care hospital use in New York State by patients infected with human immunodeficiency virus (HIV) has been based on the Statewide Planning and Research Cooperative System (SPARCS).[1] SPARCS, a mainframe data base that has been in existence since 1980, contains a discharge abstract for more than 95 percent of the acute care hospital stays in the state, exclusive of those occurring in federal and psychiatric acute care facilities. The discharge abstracts describe the hospital patient's medical condition in five diagnostic fields, including a principal diagnosis and four secondary diagnoses, using ICD-9-CM diagnostic codes.

In October 1986, new HIV-specific ICD-9-CM codes[2] became effective.[3] The new codes, 042-044, indicate the presence of certain illnesses and/or conditions occurring simultaneously with HIV infection,

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Address reprint requests to Gerald I. Kaufman, New York State Department of Health, Empire State Plaza Corning Tower, Rm. 1008, Albany, NY 12237-0045. All the authors are with the New York State Department of Health. This paper, submitted to the Journal November 20, 1989, was revised and accepted for publication June 4, 1990.

TABLE 1—HIV-Identified PCP Discharges

Year	PCP Discharges (n)	HIV-Estimated PCP[a] (n)	HIV-Estimated PCP[a] (%)[b]	HIV-Identified PCP (n)	HIV-Identified PCP ÷ HIV-Estimated PCP (%)
1987	5,725	5,631	(98)	4,891	(87)
1988	5,093	5,015[c]	(98)	4,509	(90)

Note: Data obtained from the New York State Statewide Planning and Research Cooperative System (SPARCS). Data for 1988 are 80–85 percent complete.

[a] PCP discharges minus non-HIV PCP background.

[b] HIV-estimated PCP divided by PCP discharges multiplied by 100.

[c] Background considered 82.5 percent reported (78 PCP discharges).

and another, 795.8, indicates an HIV-seropositive test result. The purpose of this study was to estimate the completeness with which admissions of HIV-infected patients were identified in the SPARCS.

Methods

We studied the completeness of HIV identification in admissions for *Pneumocystis carinii* pneumonia (PCP) in SPARCS during 1987 and 1988. PCP was extremely rare prior to the HIV epidemic[4–6]; it is now the most frequently found illness in persons with acquired immunodeficiency syndrome (AIDS).[7] Each SPARCS discharge record with a diagnosis of PCP infection (ICD-9-CM code 136.3 in the principal or secondary diagnosis fields) was considered to represent a PCP admission. Each PCP discharge with an ICD-9-CM code of 042-044 or 795.8 in the principal or secondary diagnosis fields was considered to be a PCP admission identified with HIV infection.

To establish the expected number of HIV-related PCP discharges in the given year, we subtracted the estimated number of non-HIV not-related PCP discharges from the reported number. We estimated this "background" frequency of non-HIV-related PCP discharges by counting the number of PCP discharges in 1980; we assumed such discharges to be non-HIV related. We assumed the number non-HIV-related PCP discharges had not substantially changed since 1980.

We studied the validity of the assumption that the annual number of non-HIV-related PCP discharges remained essentially unchanged in two ways. First, we compared the 1987 and 1980 PCP admissions from New York State counties that reported fewer than two AIDS cases through 1987 to the New York State AIDS Registry. If our assumption was correct, there should have been little or no difference in the number of PCP discharges by these counties in the 2 years.

Second, we reviewed the extent of HIV identification in PCP discharges in the "most sensitive" hospitals, those with the most experience with PCP in 1988. If our assumption was correct, the number of non-HIV-identified PCP discharges in such hospitals should be consistent with our 1980 background estimate.

We also investigated whether the percentage of PCP discharge abstracts with an ICD-9-CM code indicating HIV infection was associated with a hospital's experience with PCP. Hospitals were grouped into 11 different experience levels for each year studied. Linear regression using SAS[8] was performed to determine if a relation existed between PCP experience, year, and HIV coding.

Results

Completeness of HIV-Infection Identification

Non-HIV-related PCP Rates—In 1980, 94 discharges with a PCP diagnosis were reported to SPARCS. At that rate, one would expect non-HIV PCP background level of 1.6 percent in 1987 and 1988.

Eight of the 62 counties in New York State reported fewer than two AIDS cases to the New York State Registry through 1987. In 1980 each of these eight counties reported no PCP discharges. In 1987, four of the[se] eight counties reported no PCP discharges, three reported one PCP discharge, and one reported two PCP discharges.

Our review of the two hospitals with the largest number of PCP discharges (277 and 241) in 1988 revealed that HIV was associated with 97.4 percent and 99.6 percent of the PCP discharges.

PCP Discharges and HIV Codes—The results are shown in Table 1. After correction is made using the non-HIV PCP background, 98 percent of PCP discharge records in 1987 and 1988 should have been identified with an HIV diagnosis. Only 87 percent of the 1987 PCP discharge records and 90 percent of the 1988 records included an HIV diagnosis. Thus, up to 13 percent more of the 1987 and 10 percent more of the 1988 PCP discharge records should have contained an HIV-infection code.

Hospital Experience

Two linear regression models were considered. The first included year and hospital PCP volume as independent variables and percentage HIV identified as the dependent variable. A borderline significant effect was noted for year (β = 5.00, SE = 2.63, p = .07). Utilizing the

TABLE 2—HIV Associated with PCP Discharges in 1987 and 1988

PCP Discharges Per Hospital	1987 Ratio of PCP Discharges with HIV Identified to PCP Discharges (%)	1988 Ratio of PCP Discharges with HIV Identified to PCP Discharges (%)
1–10	248/389 (64)	283/368 (77)
11–20	236/29[illegible] (79)	174/202 (86)
21–30	236/267 (88)	233/269 (87)
31-40	129/169 (76)	183/210 (87)
41–50	147/181 (81)	113/131 (86)
51–60	39/52 (75)	156/162 (96)
61–70	224/267 (84)	233/254 (92)
71–80	131/148 (89)	116/142 (82)
81–90	67/83 (81)	62/81 (77)
91–100	250/289 (87)	87/99 (88)
>100	3184/3582(89)	2869/3175(90)

Note: Data obtained from the New York State Statewide Planning and Research Cooperative System (SPARCS). Data for 1988 are 80–85 percent complete.

data available from both years, a reduced model was then considered that contained only hospital PCP volume as the independent variable (Table 2). Results from that model indicated a positive association between hospital PCP volume and percentage HIV identified ($\beta = 0.93$, SE = 0.45, p = .049).

Discussion

After correcting for the non-HIV-related PCP discharges, we have found undercoding of HIV infection in discharge abstracts of PCP-identified patients in New York State's acute care facilities. The amount of undercoding was a function of the facility's volume of PCP patients.

Our estimate of HIV-related PCP discharges depended on our assumption that the number of non-HIV-related PCP discharges was essentially constant statewide between 1980 and 1988. The results from our review of PCP occurrence in counties with a low incidence of AIDS and in hospitals with a high occurrence of PCP are consistent with such an assumption. However, there may be secular trends in non-HIV-related causes of PCP that affect the background level. Review of diagnoses in the SPARCS records of non-HIV-identified PCP patients is not sufficient to determine the presence or absence of HIV infection.

The reasons for the hospital variations in reporting PCP patients as HIV infected are unknown. In the interest of confidentiality, a hospital may choose to report that a patient was not HIV-positive even though the condition was diagnosed or suspected. A lack of sufficient number of diagnostic fields in SPARCS may also be a factor, since many PCP patients are sufficiently ill with other conditions that could fill the principal and four secondary diagnosis fields of the SPARCS record. However, since the inclusion of an HIV diagnostic code would increase reimbursement substantially, HIV coding should take precedence by hospitals when its presence is known.

Another explanation is that even at this late date in the HIV epidemic, providers in New York State do not recognize that PCP patients demonstrate the most publicized HIV-related opportunistic infection. If so, other, more subtle sequelae of HIV infection may be being identified at an even lower rate, and treatment of HIV-infected patients may be delayed. Further study of the non-HIV-infected PCP patients would clearly be fruitful. In the absence of such a study, it is prudent to recognize that statewide HIV hospitalization utilization may be underestimated when SPARCS or a comparable system is used as the source. ☐

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Survival Trends of People with AIDS in Washington State

William E. Lafferty, MD, David Glidden, BA, and Sharon G. Hopkins, DVM, MPH

ABSTRACT

Survival rates of 609 cases of acquired immunodeficiency syndrome (AIDS) in Washington State diagnosed between 1982 and 1987 according to pre-1987 AIDS surveillance definition were analyzed. People with a primary diagnosis of Kaposi's sarcoma survived longer than those with *Pneumocystis carinii* pneumonia. Both groups survived longer than those with other diagnoses. Median survival increased from 11.3, to 12.5, to 20.8 months for cases diagnosed in or before 1985, during 1986, and during 1987, respectively. (*Am J Public Health* 1991; 81:215–218)

Introduction

In Seattle-King County, the major metropolitan area in Washington State, acquired immunodeficiency syndrome (AIDS) was the third leading cause of death in males aged 25–44 years in 1986. In 1987, it was second, according to the King County Department of Vital Statistics. While AIDS continues to increase in importance as a cause of mortality,[1] therapeutic advances should eventually be reflected in decreased mortality and longer survival times. Such trends have been reported from major epicenters of the epidemic, such as San Francisco,[2] and may be occurring nationwide.[3] Our study confirms that survival times of persons living

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Address reprint requests to William E. Lafferty, Office of Epidemiology and Surveillance, HIV/AIDS and Infectious Diseases, Washington State Department of Health, 1610 N.E. 150th St., Seattle, WA 98155-7224. Mr. Glidden is with the Biostatistics Department of the University of Washington. Dr. Hopkins is with the Seattle-King County Department of Public Health. This paper, submitted to the Journal September 26, 1989, was revised and accepted for publication June 21, 1990.

TABLE 1—Demographic Characteristics of AIDS Cases, by Year of Diagnosis

Year of Diagnosis	N[a]	Mean ± SD Age (Years)	% Male	% White	% Gay	% Gay/IVDU	% IVDU	Other Risk
≤1985	185	37.4±8.9	98	92	84	11	1	5
1986	181	36.7±9.3	98	90	82	6	2	6
1987	243	36.2±8.7	98	92	83	10	3	4

[a] Excludes 21 cases before 1986, 25 cases in 1986, and 25 cases in 1987 that died within 1 month of diagnosis.
Note: $p > .3$ for all comparisons of characteristics over time. IVDU = intravenous drug use.

TABLE 2—Median Survival (months) of People with AIDS in Washington State, by Disease

	Total		PCP		KS		Other Diseases	
Year of Diagnosis	N[a]	Months	N	Months	N	Months	N	Months
≤1985	185	11.3	105	12.3	34	12.4	46	7.7
1986	181	12.5	98	14.1	55	18.7	28	7.6
1987	243	20.8	137	20.8	69	—	37	14.9

[a] Excludes 21 cases before 1986, 25 cases in 1986, and 25 cases in 1987 that died within 1 month of diagnosis.

with AIDS have increased in Washington State independent of changes in the AIDS case definition.

Methods

The Washington State and Seattle-King County human immunodeficiency virus (HIV)-AIDS epidemiology units maintain [a] registry of case reports that meet the Centers for Disease Control (CDC) surveillance definition of AIDS. All resident cases were investigated. Survival analysis was performed only for patients 13 years of age or older that met the pre-1987 case definition.[4] Cases with more than one primary diagnosis were classified by hierarchy into one of three categories: *Pneumocystis carinii* pneumonia (PCP), other diagnosis without PCP, or Kaposi's sarcoma (KS) alone. Passive AIDS case reporting was estimated to be 96 percent and 88 percent complete in 1986 and 1988, respectively, according to methods previously described.[5]

In December 1988 all presumed surviving cases diagnosed on or before December 1987 were investigated. Provider contact and telephone surveillance are routine and update mortality status yearly. County vital records were reviewed for recent deaths that were either unknown to providers or may not have reached the state's vital records office. County vital records yielded 10 unreported deaths among 286 persons with AIDS who were presumed living after provider contact. Three cases lost to follow-up were censored at the last date known to be alive.

Cohorts were formed of individuals diagnosed in 1985 or before, individuals diagnosed in 1986, and individuals diagnosed in 1987. The groups were compared by means of nonparametric rank tests for censored data[6] after adjusting for primary diagnosis but not for any other prognostic variables. The effect of year and diagnosis was also estimated by fitting the Cox proportional hazards model. Hazard comparisons are relative to cases diagnosed before 1986 with KS and were made using BMDP2L.[7] Because of the limited amount of data available on any given subject, our ability to investigate and adjust for prognostic variables was limited.

The CDC case report forms record the month and year in which an individual is diagnosed and the day, month, and year of death. Approximate survival times were constructed by assuming that all individuals were diagnosed on the 15th of the month. For the generalized Wilcoxin test this is equivalent to breaking tied death times (tied if counted in months) in favor of individuals who died later in the month. These approximate death times were also used to produce Kaplan-Meier survival curves using PROC Lifetest[8] and to produce the relative hazard estimates.

Results

Through the end of 1987, 777 adult AIDS cases had been reported; 680 met the pre-1987 AIDS case definition. Seventy-one of these cases died within 1 month of diagnosis. Mean age at the time of diagnosis, racial composition, sex, and transmission category of the three temporal cohorts did not change significantly between the intervals (Table 1). Although the proportion of intravenous drug users (IVDUs) increased from 1 percent to 3

TABLE 3—Relative Hazard Table of Cohort Survival

Data	Relative Hazard	95% CI
Primary diagnosis		
KS	1.00	—
PCP	1.46	1.22, 1.91
Other diagnosis	2.20	1.63, 2.96
Year of diagnosis		
≤1985	1.00	—
1986	0.68	.545, .852
1987	0.39	.304, .497

percent from 1985 to 1987, only 12 IVDUs were included in the study. Overall, the proportion of cases diagnosed as outpatients increased from 11 percent to 20 percent between 1985 and 1987. Of definitively diagnosed cases of KS, 19 percent, 26 percent, and 68 percent were diagnosed as outpatients before 1986, in 1986, and in 1987, respectively (data not shown).

Overall, cases survived a median of 14.4 months (Tables 2 and 3). People with KS survived longer than people with PCP. Both groups survived longer than those with other diagnoses. Survival increased in cases diagnosed in more recent years. Cases of PCP and other diagnoses that were diagnosed during 1989 had longer survival times than those diagnosed in 1986 ($p < .001$, for each comparison). Most people with KS diagnosed in 1987 were still living. Median survival for inpatient cases with primary diagnoses of PCP increased from 12.4, to 15.8, to 20.8 months for the three time intervals, respectively ($p < .0001$).

Discussion

We believe our ascertainment of deaths was fairly complete. Although recent out-of-state deaths could have been missed by these methods, the impact of these cases would be 5 percent or less. In addition, since completion of this study, cases have been matched to the National Death Index. Only two deaths were detected that were not included in this analysis.

Our data confirm that trends seen in major epicenters of the AIDS epidemic are occurring elsewhere in the United States. Increases in survival for people with primary diagnoses of PCP could be related to PCP prophylaxis and zidovudine therapy. Local zidovudine studies treated approximately 52 residents who had PCP diagnosed in 1986. The lack of increase in the median survival of people with other diagnoses from 1985 to 1986 may reflect their ineligibility for zidovudine studies.

Although studies in San Francisco have shown a decrease in median survival for people with KS through 1985 and slight increases during 1986 and 1987,[2] our KS cohort diagnosed in 1986 survived longer than those diagnosed during or before 1985. The increase in outpatient diagnoses of KS from 1985 to 1987, from 19 percent to 68 percent, suggests this practice may in part be responsible for our local data. The general availability of zidovudine in March 1987 may be responsible in part for the sharp increases in survival seen in 1987. Exclusion of outpatient diagnoses for people with PCP did not reduce median survival; therefore, earlier diagnosis is an unlikely reason for prolonged survival in this cohort.

The impact of prolonged survival of people living with AIDS on the health care system needs evaluation. Although CDC case projections for 1990 and beyond have been reduced in cumulative numbers by 16 percent through 1993,[9] increased survival may neutralize the effect of fewer new cases on the health care system. Recent federal registration (H.R. 2310) expanding insurance continuation from 18 to 29 months may make health care more accessible to a larger number of people living with AIDS who survive longer than 18 months. However as the use of zidovudine and other antiviral drugs prolongs survival, health care programs need to provide long-term access to people who are otherwise uninsurable. ☐

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New from NCHS

Long-term Care for the Functionally Dependent Elderly

Introduction

Greater attention has focused recently on financing of long-term care, due both to the growing burden to society and individuals and to the increasing size of the elderly population at risk of needing long-term care. As a result, insurance companies underwriting private long-term care insurance policies, legislators, and other policy makers need estimates of the population most at risk of needing long-term care. A recent report from the National Center for Health Statistics (NCHS) should help such policy makers plan for long-term care services.[1]

The report was based on combined data from the Supplement on Aging (SOA) to the 1984 National Health Interview Survey (NHIS) and the 1985 National Nursing Home Survey (NNHS). In brief, the NHIS is a continuing national survey of the health of people residing in households in the United States. The SOA was added in 1984 to obtain information about older persons that had not been available previously from a national survey. All persons age 65 years and over and half of those ages 55–64 in responding NHIS households were included in the SOA sample; 97 percent of those eligible participated in the SOA (16,148 people) of whom 92 percent responded for themselves. Descriptions of the NHIS and SOA have been published.[2,3]

Data on nursing home residents are from the 1985 NNHS, a national probability sample of nursing homes, their residents, discharges, and staff. Homes included in the NNHS were nursing and related care homes in the conterminous United States, without regard to the level of care, whether they participated in Medicare or Medicaid, or whether they were licensed. Board and care homes and residential care homes were excluded.

Nursing home resident estimates are based on the sample of 4,650 residents ages 65 years and over from the 1,079 nursing homes participating in the survey. Residents included in the sample were those on the nursing home's roster as of the night before survey. Data were collected by interviewing knowledgeable nursing home staff members, who referred to the residents' medical records when necessary. The response rate for the resident sample was 97 percent.[4]

The report profiles long-term care use by functionally dependent people age 65 years and over, where functionally dependent is defined as receiving human assistance with at least one activity of daily living (ADL) or instrumental activity of daily living (IADL). Persons who did not perform or were unable to perform the activity were also considered dependent. ADLs are the basic personal functions of bathing, dressing, eating, getting in or out of a bed or chair, mobility, using the toilet, and continence. IADLs are everyday activities that enable an individual to live independently in the community: preparing meals, shopping, managing money, using the telephone, doing light or heavy housework, and getting around outside the home.

Criteria used to define dependence in the ADLs and IADLs for noninstitutionalized people were largely based on those used in an earlier report on the physical functioning of the aged.[5]

Noninstitutionalized Elderly

Dependence in the ADLs for noninstitutionalized people was defined as: 1) having difficulty performing the activity because of a chronic condition and *receiving* help of another person in performing the activity; or 2) being unable to perform the activity without special equipment and not having that equipment. Dependence in the IADLs for this population was defined as having difficulty or inability to perform the activity without assistance because of a health problem.

Institutionalized Elderly

On the other hand, dependence in the ADLs for nursing home residents was defined as requiring (needing) assistance of another person or special equipment to perform the activity. Dependence in the IADLs for nursing home residents was defined as receiving personal help or supervision to perform the activity. For the purposes of this report, it was assumed that the effect of differences in question wording between the SOA and the NNHS ("receiving" versus "needing" help) was minimal because residents in an institutional environment are likely to receive assistance if they need it. It was also assumed that all nursing home residents were dependent in two IADL activities (preparing meals and doing heavy housework) routinely provided in nursing homes, even though not all residents were dependent in these activities.

Results

Based on combined estimates of the institutionalized and noninstitutionalized population of the United States, 29 percent of people age 65 years and over in

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Address reprint requests to Esther Hing, Division of Health Care Statistics, National Center for Health Statistics, 6525 Belcrest Road, Room 950, Hyattsville, MD 20782.

TABLE 1—Number and percent of functionally dependent people age 65 years and over by level of ADL or IADL dependency and type of long-term care received, United States, 1984–85

Level of ADL or IADL Dependency and Type of Long-term Care Received	Number	Percent of Persons aged 65 Years and over	Percent of Functionally Dependent
All persons 65 years and over	27,751,300	100.0	—
Level of ADL or IADL dependency			
All functionally dependent elderly	8,051,100	29.0	100.0
5–7 ADL dependencies	1,373,800	5.0	17.1
3–4 ADL dependencies	728,000	2.6	9.0
1–2 ADL dependencies	3,968,600	14.3	49.3
IADL dependencies only	1,960,700	7.1	24.6
Type of long-term care received by functionally dependent elderly			
All functionally dependent elderly	8,051,100	29.0	100.0
Nursing home care	1,318,300	4.8	16.4
Formal home care[a]	1,662,500	6.0	20.6
Informal home care only	3,380,400	12.2	42.0
No long-term care	1,689,800	6.1	21.0

a) Includes persons receiving both formal and informal home care services.
ADL = activities of daily living
IADL = instrumental activities of daily living.

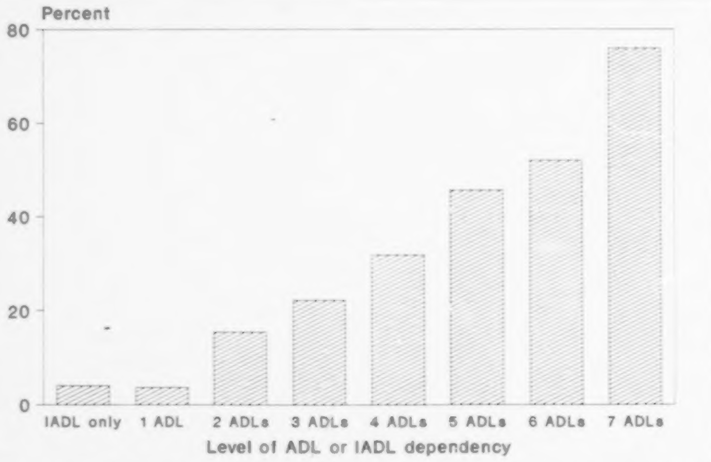

FIGURE 1—Percent of functionally dependent persons age 65 years and over residing in nursing homes by level of ADL or IADL dependency: United States, 1984–85

1984–85, or 8 million persons, were functionally dependent and at risk of needing long-term care services. About three-fourths of this population had only slight limitations in these functions; 49 percent were dependent in only one or two ADLs and 25 percent were dependent in only the IADLs. About 9 percent were moderately dependent, receiving assistance in three to four ADLs, while 17 percent were severely dependent, receiving assistance in nearly all ADLs and IADLs.

Although long-term care has traditionally meant nursing homes, it also includes the services of home health agencies, visiting nurses, adult day care centers and other community-based services, as well as care provided informally by family and friends. This report examined three types of long-term care: nursing home care, formal home care, and informal home care. Formal home care services were defined as receipt of paid-for personal assistance in performing at least one ADL or IADL, while help provided without compensation was considered informal home care. In 1984–85, 79 percent of the functionally dependent elderly population received long-term care services (Table 1). Informal home care by family and friends was the most commonly received type of long-term care (42 percent). About 21 percent received formal home care services, often in combination with informal home care. About 16 percent of the functionally dependent elderly resided in nursing homes. The percent of functionally dependent elderly persons receiving these three types of long-term care is much higher than that for the general elderly population. In 1984–85, the percent of elderly receiving informal home care was 12 percent, while 6 percent received at least some formal home care service. Only 5 percent of the elderly resided in nursing homes (Table 1).

The degree of dependency in the ADLs and IADLs was associated with the type of long-term care used by functionally dependent persons. Home care use was greater among those with only IADL limitations or four or fewer ADL limitations; nursing home use was greatest among those dependent in five to seven ADLs. There was a nearly linear relationship between the percent of functionally dependent elderly residing in nursing homes and the number of ADL dependencies. About 76 percent of the elderly dependent in all ADL functions were in nursing homes compared with 4 percent of the elderly dependent in only one ADL function (Figure 1). This finding, however, is affected by the fact that many nursing homes have policies that do not permit a resident to perform the ADLs or IADLs without assistance.

Because functional dependency in ADL and IADL activities is higher at older ages (from 20 percent among those ages 65–74 years to 66 percent among those ages 85 years and over), overall use of long-term care, and use of nursing homes in particular, was also higher at older ages. The percent of functionally dependent elderly using long-term care increased from 71 percent among those ages 65–74 years to 93 percent of those ages 85 years and over; nursing home use increased from 6 percent among those ages 65–74 years to 36 percent among those age 85 years and over. The average age of functionally dependent elderly in nursing homes was 83 years, while that of their counterparts using any formal or only informal home care services was 78 and 77 years, respectively. Besides age, other

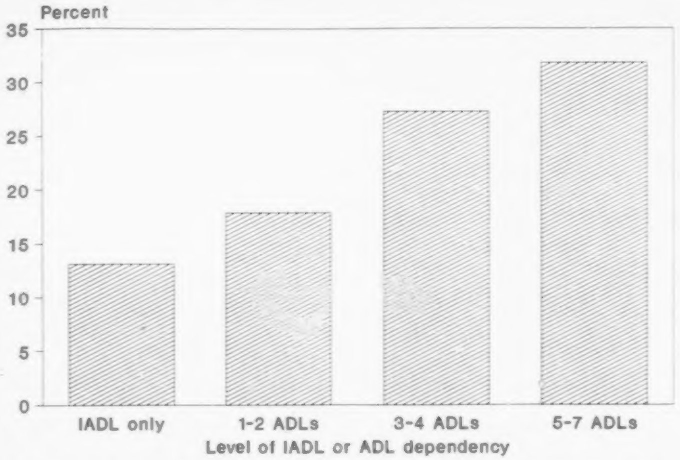

FIGURE 2—Percent of functionally dependent home care recipients receiving both formal and informal home care services: United States, 1984–85

factors found to be significantly associated with nursing home use were being female, White race, being unmarried, and having income below the poverty threshold.

Home care services (formal or informal) were used by a younger, less dependent population. Availability of an informal support network of family and friends appears to be associated with receipt of informal care; among the functionally dependent elderly, use of informal home care services was greater for males, for persons living with a spouse or others, and for married persons; use of any formal home care services was greater for females, for persons living alone, and for unmarried persons. The greater use of informal home care services by Black functionally dependent elderly coincides with their greater likelihood of living with spouse or others (potential caregivers). Functionally dependent elderly persons with income above the poverty threshold were more likely to use both formal and informal home care services.

Use of formal home care services, however, was also dependent on the level of ADL or IADL limitations. Formal home care services were used to augment informal care provided by family or friends more often among the severely dependent elderly than among those with few limitations. Thirty-two percent of home care recipients dependent in five to seven ADLs, and 27 percent of the elderly home care recipients dependent in three to four ADLs relied on both formal and informal caregivers, compared with 13 percent of home care recipients dependent in only the IADLs (Figure 2). Persons living alone were four to five times more likely to use only formal home care services than those living with spouse or others; persons living with spouse or others were more likely to receive only informal home care services.

Discussion

While this report presents a more complete picture of long-term care use by the functionally dependent elderly population than available by looking at only the nursing home or the community-dwelling population, the data have limitations in interpretation. First, there are reporting differences in the SOA and NNHS data; the surveys had different respondents (self-report versus nursing home staff), and the question wording was not always identical. Second, the two surveys were fielded in different years. This problem, however, should be negligible since the proportion of the noninstitutionalized population that was functionally dependent did not change between 1984 and 1986. Further research is possible by accessing public use data tapes for the 1984 SOA and the 1985 NNHS. The 1984 SOA can be obtained from the Division of Health Interview Statistics, National Center for Health Statistics in Hyattsville, Maryland. The 1985 NNHS can be obtained from the National Technical Information Service in Springfield, Virginia. Both data tapes are also available from the National Archive of Computerized Data on Aging (NACDA) from the Inter-university Consortium for Political and Social Research at the University of Michigan in Ann Arbor, Michigan. ☐

Esther Hing, BS
Barbara Bloom, MPA

Copies of the report, *Long-term Care for the Functionally Dependent Elderly*, by Hing and Bloom, *Vital Health Stat* 13 (104) can be purchased from the US Government Printing Office, Washington, DC 20402. Order stock no. 017-022-01116-6. The price is $3.00.

For more information about this and other reports and data from the National Center for Health Statistics, write to the Scientific and Technical Information Branch, NCHS, 6525 Belcrest Road, Hyattsville, MD 20782, or call (301) 436-8500.

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Impact of the Minnesota Parental Notification Law on Abortion and Birth

James L. Rogers, PhD, Robert F. Boruch, PhD, George B. Stoms, BA, and Dorothy DeMoya, DNSc

ABSTRACT

Background. The impact of the Minnesota Parental Notification Law on abortion and birth was examined.

Methods. Using linear models, outcome parameters were compared before and after enactment of the law. Time by age group interactions also were examined.

Results. The pre-enactment to post-enactment change in the Minnesota abortion rate reflected a greater decline for minors (≤17 years old) than for 18–19 year-olds (who were not under the law). An increase in abortion rate occurred for women ages 20–44. The law appeared to have had no impact on birth rate in minors. Following the enactment of the law, the rate of early abortions (≤12 weeks) declined among minors more than the rate of late abortions (>12 weeks). This resulted in a pre-enactment to post-enactment increase in the ratio of late-to-early abortions among minors.

Conclusions. These data suggest that parental notification facilitated pregnancy avoidance in 15–17 year-old Minnesota women. Abortion rates declined unexpectedly while birth rates continued to decline in accordance with a long-term trend. (*Am J Public Health* 1991;81:294–298)

Introduction

Laws requiring parental consent or parental notification prior to legal induced abortion for minor women, collectively called parental involvement laws, exist or have been proposed in numerous states. As of July 1990, laws in the United States requiring parental consent were in effect in Alabama, Indiana, Louisiana, Massachusetts, Missouri, North Dakota, and Rhode Island. Laws requiring parental notice were in effect in Arkansas, Idaho, Utah, and West Virginia; and parental involvement statutes were under challenge in Arizona, California, Georgia, Illinois, Kentucky, Mississippi, Nevada, Pennsylvania, and Tennessee. National attention focused on these laws when statutes from Minnesota and Ohio were heard by the US Supreme Court during its October 1989 term resulting in a decision largely supporting these laws. The present paper concerns the Minnesota law, enacted in August 1981 and enjoined in March 1986. This law required a minor woman to notify both parents at least 48 hours prior to an abortion or else seek court approval.

Few empirical studies have evaluated the impact of parental involvement statutes on minor women. Cartoof and Klerman[1] determined that abortions to minors in Massachusetts declined dramatically (43 percent) following the enactment of a parental consent law. However, during this time an approximately equal number of women migrated to surrounding states to obtain abortions. Blum[2] found that under parental notification in Minnesota, communication with parents about a minor's planned abortion occurred more often than had been reported by Clary[3] in a Minneapolis/St. Paul study predating the law. But Blum found that patterns of communication differed little from those among teenagers simultaneously surveyed in the neighboring state of Wisconsin (without such a law).

Common negative claims about parental involvement laws are that they force minors to leave the state to obtain abortions (as in Massachusetts), and that they result in increased birth rates, late abortions and medical complications. These effects are presumably related to a minor's reluctance to discuss her pregnancy with parents.[4] Positive claims about these laws are that they promote responsibility (by encouraging teenagers to "think before they act"), foster parent-child communication, facilitate mature decision making, and may reveal medical history information that would otherwise remain unknown to the physician.[5,6]

Empirical evaluation of assertions like these will necessitate multiple studies under a variety of circumstances and localities. The Cartoof and Klerman study[1] was conducted in Massachusetts, located in close proximity to states without parental involvement laws. This made it possible for minors to avoid the law altogether by crossing state lines. In Minnesota, the distance from out-of-state abortion facilities appears to have worked against mi-

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Address reprint requests to James L. Rogers, PhD, Professor, Department of Psychology, Wheaton College, Wheaton, IL 60187. He is also Senior Research Associate, Department of Medicine, Northwestern University Medical School. Dr. Boruch is University Trustee Professor, Graduate School of Education, and Department of Statistics, Wharton School, University of Pennsylvania; Mr. Stoms is with the Department of Psychology, Wheaton College; Dr. DeMoya is Consultant/Therapist, Family Resources, Philadelphia, PA. This paper, submitted to the Journal July 17, 1989, was revised and accepted for publication October 10, 1990.

gration. Blum determined that "[i]n counterdistinction to the Massachusetts data, there is little evidence to indicate large numbers of Minnesota youths are leaving the state for abortion (data available on request to author)."[2] It cannot be assumed that findings characterized by one set of background factors, such as proximity to out-of-state abortion facilities, will generalize to other settings.

In this study, the statewide impact of the Minnesota Parental Notification Law upon the incidence rate of abortion and birth, as well as upon the ratio of abortions to births and the ratio of early to late abortions, is examined.

Methods

Data

Abortion and birth incidence data were provided by the Minnesota Center for Health Statistics (MCHS). The data exclude all observations of unknown age and are restricted to residents of Minnesota. Live births to Minnesota residents are included regardless of whether the birth occurred inside or outside of Minnesota. Induced abortions reflect only those occurring in Minnesota.

Population estimates by age and gender are provided by the Minnesota Center for Health Statistics that computed them using a modified version of the cohort-component method for all years following the 1980 census.[7]

Throughout this report "birth(s)" and "abortion(s)" will refer to live birth(s) and induced abortions(s), respectively.

Outcome Measurements

The report utilizes six outcome measurements: four rates and two ratios.

- The abortion rate, the late abortion rate (>12 weeks), the early abortion rate (≤12 weeks) and the birth rate refer to the number of reported abortions (or births) in one year divided by the population estimate of females, in thousands, for that same year.
- The abortion-to-birth ratio refers to the number of abortions in a year divided by the number of births. Alternatively, this may be thought of as the abortion rate divided by the birth rate for a given year.
- The late-to-early abortion ratio refers to the number of late abortions in a year divided by the number of early abortions. Again, this may be thought of as the late abortion rate divided by the early abortion rate for a given year.

Measures of Effect

Each rate and ratio was examined using a linear model.[8,9] Serving as a dependent variable, the rate (or ratio) was modeled as a function of age category (≤17, 18–19 or 20–44 years old), the year of occurrence (1975 through 1987), and the age by year interaction.

First, each model was employed to determine whether a given rate (or ratio) three years before and four years after enactment of the Minnesota Parental Notification Law differed within each age category. Because the modeling was performed in the log scale, the pre-enactment (1978 to 1980) and post-enactment (1982 to 1985) values represent the geometric mean of the individual values comprising the pre-enactment and post-enactment periods. (The antilog of the arithmetic mean of log values corresponds to the geometric mean of the same measurement in the original scale. That is, anti-ln $[\ln a + \ln b]/2 = \sqrt{[ab]}$.)

Second, three additional contrasts were constructed to detect the presence of any age group by time interaction that might exist for a given rate or ratio. These contrasts reflect whether the pre-enactment to post-enactment *change* was different among minors than among 18–19 year-olds, or 20–44 year-olds, or among 18–19 year-olds than women 20–44 years old. It was assumed that a change due to the law, rather than to general factors operating in all age groups, would be most pronounced among women 17 years of age or younger; less evident among 18 and 19 year-old women who would have recently been, but would not presently be under the law (pregnancy at age 17 may mean birth at age 18); and least present among older women not subject to the law for at least two years.

Models

The mechanics underlying the linear models[8,9] used to construct the six contrasts described above were as follows. The model parameters, representing age category (two parameters capturing three age classifications), year (12 parameters capturing 13 years), and the age by year interaction (24 parameters reflecting the cross-product of age and year), were regressed against the ***natural log*** of the rate or ratio under question. Rows of each model's design matrix were combined to form the six contrasts. When the abortion rate, late abortion rate, early abortion rate, or birth rate served as the dependent variable, weighted least squares estimates and asymptotic variances for the estimates were obtained. When the abortion-to-birth ratio or late-to-early abortion ratio served as the dependent variable, maximum likelihood was used to obtain estimates and asymptotic variances. PROC CATMOD of Version 6.03 of the Statistical Analysis Software (SAS)[10] was employed to fit the models.

For ease of interpretation, the authors elected to display each contrast effect as a quotient (contrast ratio) in the original scale rather than a difference in the log scale. For any given contrast, this means that rather than presenting in tables the difference between two natural log values, it is the antilog of this difference that has been presented. It is evident that the difference between two identical log values will be "zero" while the corresponding contrast ratio will be unity (one). That is, $(\ln A) - (\ln A) = 0$ implies that the antilog is unity. Thus, contrast ratios equal to unity imply equivalence between the contrasted values.

Results

Table 1 contains the outcome measures examined in this study. For each outcome measure, Table 2 contains the contrast ratios that compare the pre-enactment and post-enactment periods. Contrast ratios greater than unity imply an increase in the outcome measure (abortion rate, birth rate, etc.) after enactment of the law and contrast ratios less than unity imply a decrease. Similarly, Table 3 contains ratios that reflect the age by time interactions. Here, a contrast ratio less than unity indicates a greater pre-enactment to post-enactment decline in the younger age group of the two being compared; a contrast ratio greater than unity indicates a greater increase.

Abortion Rate

Deviations from unity for the contrast ratios that compare pre-enactment and post-enactment periods (Table 2) are substantial in all age groups. Whereas the yearly abortion rates after the law's enactment increased for women 20–44 years old (who were substantially removed from its impact), abortion rates declined in both 15–17 and 18–19 year-olds during this same period. The pre-enactment to post-enactment decline was substantially greater for 15–17 than 18–19 year-old women, and for 18–19 year-old women than 20–44 year-old women (Table 3).

TABLE 1—Outcome Measures and Population Estimates for Minnesota Women, 1975 to 1987*

Outcome Measure	Age (years)	1975	1976	1977	1978	1979	1980	1981	1982	1983	1984	1985	1986	1987
Abortion Rate†	15–17	12.38	16.50	18.60	18.25	19.24	19.57	16.06	14.25	12.80	13.03	14.54	14.42	15.45
	18–19	20.47	28.92	31.27	35.71	38.72	40.26	38.37	36.45	33.08	35.05	34.07	31.89	30.83
	20–44	7.32	9.55	11.48	12.36	13.41	14.13	14.00	13.96	13.11	14.21	14.48	14.29	14.17
Birth Rate†	15–17	20.94	19.64	19.62	17.80	17.71	17.48	17.36	16.54	14.56	16.00	15.01	15.52	15.03
	18–19	56.11	52.96	55.10	55.60	57.00	59.48	59.33	56.57	48.76	48.85	47.18	42.65	43.68
	20–44	74.02	72.48	75.79	76.14	78.05	79.80	79.43	78.13	75.18	75.57	75.92	74.00	72.67
Abortions/ Births	15–17	0.59	0.84	0.95	1.03	1.09	1.12	0.93	0.86	0.88	0.81	0.97	0.93	1.03
	18–19	0.36	0.55	0.57	0.64	0.68	0.68	0.65	0.64	0.68	0.72	0.72	0.75	0.71
	20–44	0.10	0.13	0.15	0.16	0.17	0.18	0.18	0.18	0.17	0.19	0.19	0.19	0.19
Early Abortion Rate†	15–17	10.22	12.81	14.73	14.97	15.73	15.34	12.93	11.37	9.86	9.68	11.24	11.38	12.65
	18–19	17.81	24.01	25.89	30.46	33.31	33.56	32.58	30.92	27.63	28.42	28.13	26.04	26.23
	20–44	6.62	8.45	10.23	11.24	12.19	12.71	12.74	12.65	11.87	12.69	13.04	12.85	12.84
Late Abortion Rate†	15–17	2.16	3.69	3.88	3.28	3.51	4.23	3.13	2.89	2.94	3.36	3.30	3.04	2.80
	18–19	2.65	4.91	5.39	5.25	5.41	6.69	5.78	5.53	5.44	6.63	5.94	5.85	4.56
	20–44	0.70	1.10	1.25	1.11	1.23	1.42	1.27	1.31	1.24	1.53	1.44	1.44	1.31
Late/Early Abortions	15–17	0.21	0.29	0.26	0.22	0.22	0.28	0.24	0.25	0.30	0.35	0.29	0.27	0.22
	18–19	0.15	0.20	0.21	0.17	0.16	0.20	0.16	0.18	0.20	0.23	0.21	0.22	0.17
	20–44	0.11	0.13	0.12	0.10	0.10	0.11	0.10	0.10	0.10	0.12	0.11	0.11	0.10
Population	15–17	115684	117102	116317	115722	115262	113600	108143	103981	104371	100131	100912	101172	101846
	18–19	85889	86826	86110	85525	85047	83964	79863	76786	77004	73784	74295	74375	74788
	20–44	662309	683069	700327	722162	747058	761209	779081	797138	799912	811693	819042	821954	828167

*Raw data provided by the Minnesota Center for Health Statistics.

†Abortion, birth, early abortion and late abortion rates are expressed as the number of abortions or births per 1000 women.

NOTE: Early abortions: ≤12 weeks; Late abortions: >12 weeks.

Birth Rate

Birth rates decreased in all age categories following enactment of the law (Table 2). However, the decline was most pronounced in 15–17 and 18–19 year-old women. Table 3 reveals that the pre-enactment to post-enactment change among 15–17 and 18–19 year-old women was similar, with both age groups evidencing a substantially greater decline than found among women ages 20–44.

Ratio of Abortions to Births

A marked drop in the abortion-to-birth ratio occurred after the law in 15–17 year-old women when compared to both 18–19 year-old women and 20–44 year-old women (see Tables 2 and 3). In Figure 1, the abortion rate and birth rate are plotted separately for 15–17 year-old women along with the abortion-to-birth ratio (abortion rate/birth rate) in order to examine the relative importance of abortions and births to the markedly declining abortion-to-birth ratio in this age group. It is evident that birth rates continue a modest and nearly linear decline, apparently unaffected by the law ($r = -0.89$ between birth rate and year). On the other hand, the abortion rate falls dramatically after the enactment of the law in August 1981. Together, these facts indicate that the drop in the 15–17 year-old abortion-to-birth ratio is due to a disproportionately greater decrease in the abortion rate (numerator).

Early and Late Abortions

The early abortion rate closely tracks the overall abortion rate (Tables 2 and 3). The pre-enactment to post-enactment late abortion rate substantially declines for women of 15–17 years, increases for women of 20–44 years, and remains nearly constant for women of 18–19 years (Table 2). The pre-enactment to post-enactment change in the late abortion rate, when compared between age groups, evidences a greater decline in late abortions for 15–17 than for either 18–19 or 20–44 year-old women (Table 3).

The late-to-early abortion ratio increased after the enactment of the law in all age groups (Table 2). However, the increase was greater among 15–17 year-old women than 20–44 year-old women (Table 3). Figure 2 reveals that a steep decline in early abortions, not an increase in late abortions, accounts for the increased late-to-early abortion ratio in 15–17 year-old women.

Discussion

Data presented in this study are compatible with the hypothesis that, initially, parental notification facilitated pregnancy avoidance in 15–17 year-old Minnesota women. Abortion rates fell markedly in this age group relative to older women. Birth rates also fell, but only in keeping with a long-term trend established before enactment of the law. One possibility is that when minor women are restricted from abortion without notifying parents or seeking court approval, and are geographically prohibited from easy access to out-of-state abortions,[2] they are more likely to take measures to avoid pregnancy.

Although the data are compatible with this hypothesis, other explanations are possible. For example, a growing concern over human immunodeficiency virus infection, and/or awareness and availability of birth control may explain in part or in full these findings. However, the abrupt nature of the change in abortion rate, a phenomenon found also in Massachusetts by Cartoof and Klerman,[1] makes these rival hypotheses less tenable. In any event, the data argue against Clary's[3] concern that more minors might carry pregnancies to term as an indirect effect of the parental notification law. If such were the case, it seems unlikely that birth rates would have continued to decline in 15–17 year-olds along the linear trend line established prior to the law, or that the decline in birth rates would be nearly identical between 15–17 and 18–19 year-old women.

The pre-enactment to post-enactment increase in the proportion of late (>12 weeks) to early (≤12 weeks) abortions was greater for 15–17 than for 20–44 year-old women. At least two hypotheses may explain this finding. First, the law may have been more successful in preventing pregnancy among minors who would have had early abortions than among minors who would have had late abortions. A second possibility is that the law caused delays for a greater percentage of a declining number of minors seeking abortions. Regardless, the claim that the law caused more minors to obtain late abortions is unsubstantiated. In fact, the reverse is true. For ages 15–17 the number of late abortions per 1,000 women decreased following the enactment of the law. Therefore, an increased medical hazard due to a rising number of late abortions was not realized.

In this paper no effort has been made to confront the philosophical and legal issues surrounding parental involvement laws. Rather, the authors have pursued a limited task, that of empirical evaluation within a framework of defined outcome parameters. This study is consistent with the hypothesis that conception among minor women may be reduced immediately following enactment of parental notification legislation when migratory abortion across state lines is not a viable alternative. However, generalizations to other states must be made cautiously, as Minnesota is a unique state with a low minority population and a low pregnancy rate even before the parental notice legislation. The authors emphasize that replication in states other than Minnesota will be required to sustain the hypothesis. ☐

TABLE 2—Contrasts Between Pre- and Post-Enactment Periods[*]

Outcome Measures	Age (years)	Pre-Enactment[**]	Post-Enactment[†]	Contrast Ratio (Post/Pre) With 95% CI
Abortion Rate[††]	15–17	19.012	13.635	0.717 (0.692, 0.743)
	18–19	38.181	34.638	0.907 (0.883, 0.932)
	20–44	13.280	13.931	1.049 (1.034, 1.064)
Birth Rate[††]	15–17	17.663	15.510	0.878 (0.848, 0.909)
	18–19	57.338	50.213	0.876 (0.857, 0.895)
	20–44	77.982	76.191	0.977 (0.971, 0.983)
Abortions/Births	15–17	1.076	0.879	0.817 (0.777, 0.859)
	18–19	0.666	0.690	1.036 (1.000, 1.074)
	20–44	0.170	0.183	1.074 (1.057, 1.091)
Early Abortion Rate[††]	15–17	15.343	10.507	0.685 (0.658, 0.713)
	18–19	32.413	28.749	0.887 (0.861, 0.914)
	20–44	12.032	12.554	1.043 (1.027, 1.059)
Late Abortion Rate[††]	15–17	3.653	3.114	0.852 (0.788, 0.921)
	18–19	5.750	5.867	1.020 (0.952, 1.093)
	20–44	1.247	1.375	1.103 (1.052, 1.157)
Late/Early Abortions	15–17	0.238	0.296	1.245 (1.140, 1.359)
	18–19	0.177	0.204	1.150 (1.067, 1.241)
	20–44	0.104	0.110	1.058 (1.006, 1.112)

[*] Raw data provided by the Minnesota Center for Health Statistics.

[**] Geometric mean, years 1978–80, Table 1.

[†] Geometric mean, years 1982–85, Table 1.

[††] Abortion, birth, early abortion and late abortion rates are expressed as the number of abortions or births per 1000 women.

NOTES: 1) Early abortions: ≤12 weeks; Late abortions: >12 weeks.
2) Pre-enactment (1978–80) to post-enactment (1982–85) means are compared (post/pre) in the form of contrast ratios. A contrast ratio of one implies no pre-enactment to post-enactment change.

TABLE 3—Age by Time Interactions[*]

Outcome Measures	Age Group Comparison	Post- / Pre-enactment Ratios[**]		Contrast Ratio (younger/older) with 95% CI
		Younger	Older	
Abortion Rate	15–17 vs 18–19	0.717	0.907	0.791 (0.756, 0.827)
	15–17 vs 20–44	0.717	1.049	0.684 (0.658, 0.710)
	18–19 vs 20–44	0.907	1.049	0.865 (0.839, 0.892)
Birth Rate	15–17 vs 18–19	0.878	0.876	1.003 (0.962, 1.045)
	15–17 vs 20–44	0.878	0.977	0.899 (0.867, 0.931)
	18–19 vs 20–44	0.876	0.977	0.896 (0.876, 0.917)
Abortions/Births	15–17 vs 18–19	0.817	1.036	0.788 (0.741, 0.839)
	15–17 vs 20–44	0.817	1.074	0.761 (0.722, 0.802)
	18–19 vs 20–44	1.036	1.074	0.965 (0.928, 1.003)
Early Abortions	15–17 vs 18–19	0.685	0.887	0.772 (0.735, 0.812)
	15–17 vs 20–44	0.685	1.043	0.656 (0.629, 0.685)
	18–19 vs 20–44	0.887	1.043	0.850 (0.822, 0.879)
Late Abortions	15–17 vs 18–19	0.852	1.020	0.835 (0.753, 0.927)
	15–17 vs 20–44	0.852	1.103	0.772 (0.705, 0.846)
	18–19 vs 20–44	1.020	1.103	0.925 (0.850, 1.006)
Late/Early Abortions	15–17 vs 18–19	1.245	1.150	1.082 (0.963, 1.215)
	15–17 vs 20–44	1.245	1.058	1.177 (1.064, 1.302)
	18–19 vs 20–44	1.150	1.058	1.088 (0.994, 1.191)

[*] Raw data provided by the Minnesota Center for Health Statistics.

[**] Post- / pre-enactment ratios are from Table 2.

NOTES: 1) Early abortions: ≤12 weeks; Late abortions: >12 weeks.
2) Post-enactment to pre-enactment ratios (Table 2) are compared across age groups (younger/older) to examine age by time interactions. A contrast ratio of "one" implies equivalent post- / pre-enactment ratios for both age groups (no interaction).

Acknowledgments

The authors wish to thank the Minnesota Center for Health Statistics, particularly James Wigginton and Carol Vargas for their consid-

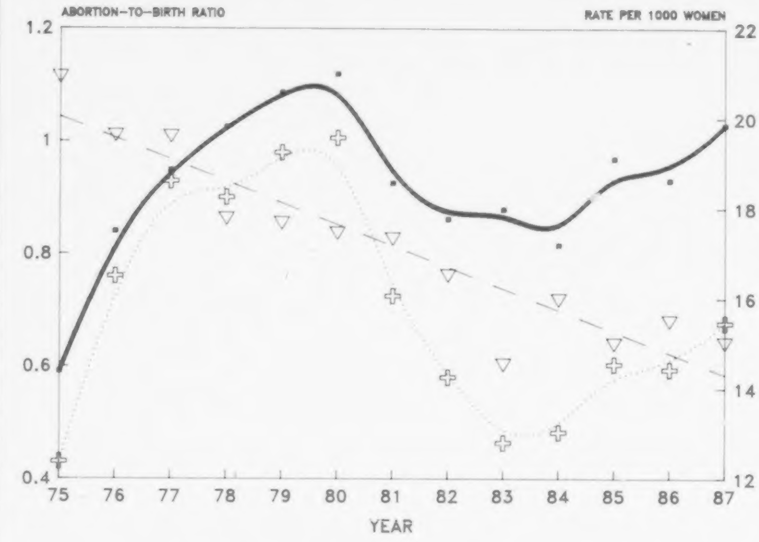

FIGURE 1—Abortion and Birth Rates per 1000 15–17 Year-Old Women and Abortion/Birth Rate Ratio, 1975–87
—Abortion/Birth Ratio
···+··· Abortion Rate
—▽— Birth Rate

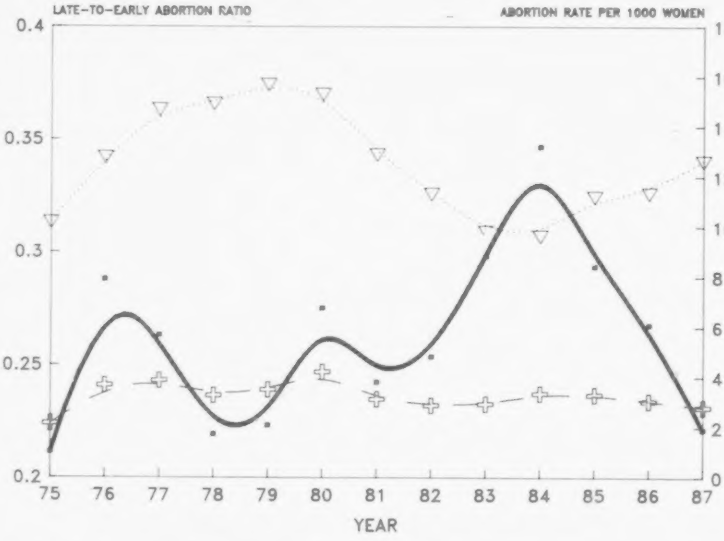

FIGURE 2—Early and Late Abortion Rates per 1000 15–17 Year-old Women and Late/Early Abortion Rate Ratio 1975–87
—Late/Early Ratio
—+— Early Abortion Rate
···▽··· Late Abortion Rate

erable work in providing these data in a form suitable for the current analysis. The authors also thank Americans United for Life for purchasing the data from the Minnesota Center for Health Statistics and providing it to the authors. Preparation of this study was supported in part by an Aldeen Grant from Wheaton College.

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Psychosocial Factors in Maternal Phenylketonuria: Prevention of Unplanned Pregnancies

Susan E. Waisbren, PhD, Shoshana Shiloh, PhD, Paula St. James, MA, and Harvey L. Levy, MD

ABSTRACT

Background. Women with phenylketonuria (PKU) not treated prior to conception can have a pregnancy that results in serious fetal damage. In this report, factors associated with preventing unplanned (and hence late treated) pregnancies are described.

Methods. Subjects included 60 phenylketonuric women and two comparison groups composed of female acquaintances and diabetic women. All were interviewed and administered tests of intelligence, general well-being, knowledge, and personality.

Results. Thirty-five percent of the sexually active women with PKU used contraception only sporadically. The variables that best predicted reported frequency of birth control use were the extent to which women felt social support to use contraception ($r = .64$) along with positive attitudes about birth control ($r = .66$) and knowledge of family planning ($r = .43$). For the comparison groups, a different pattern of variables predicted contraceptive use, with locus of control figuring most prominently for the diabetics ($r = .39$) and social support for birth control being most important for the acquaintances ($r = .46$).

Conclusions. As more girls with PKU enter childbearing ages, there will be an increased need for specific programs that address psychosocial factors in maternal PKU. (*Am J Public Health* 1991;81:299–304)

Introduction

Maternal phenylketonuria (PKU) as a cause of mental retardation and birth defects is a new phenomenon. In the past, most young women with PKU (an inborn error of phenylalanine metabolism) were mentally retarded and few bore children. Now, most such women are mentally normal, having been born since newborn screening was initiated in the 1960s and treated from early infancy with a phenylalanine-restricted diet.[1] In the United States there may be as many as 4,000 women of childbearing age with PKU,[2] the majority of whom are at risk for pregnancies that can result in abnormal children. For those with the most severe and most common form of PKU, the frequencies of complications in offspring from untreated pregnancies are 92 percent for mental retardation, 73 percent for microcephaly, 12 percent for congenital heart disease, and 40 percent for low birth weight.[3] In fact, it is feared that in the next generation, the consequences of maternal PKU could replace the number of individuals with PKU in whom mental retardation has been prevented by newborn screening.[4, 5]

Dietary treatment will protect the fetus but only under the proper circumstances. Treatment begun prior to conception and continuing throughout pregnancy appears to offer maximum protection, whereas treatment even in the first trimester may not prevent some of the adverse effects of maternal PKU.[6,7] Since many young women with PKU discontinued the special diet during childhood,[8] the strategy for preventing birth defects from maternal PKU must include careful family planning so that diet can be initiated before conception.

Tracking adolescents and young adult women with PKU so that they can be advised of the high risk to their pregnancies and the possibilities of treatment is difficult, but with much effort can be done.[2] In Quebec, there is a registry that tracks individuals with PKU through adulthood.[9] The results of other tracking efforts, however, indicate that women with PKU usually seek treatment after they are pregnant rather than before.[7,10,11] The reasons for this are unknown.

The overall purpose of our study is to identify these reasons. We will describe how women with PKU face questions of sexuality and childbearing in order to determine the factors related to preventing unplanned pregnancies in maternal PKU.

We describe here the findings from the initial interviews of all subjects who were not planning a pregnancy and were not pregnant when the study began in 1985.

Methods

Subjects

Young women in New England known to have hyperphenylalaninemia, ages 16 to 35 years, were invited to participate in a five-year longitudinal study.

Address reprint requests to Susan E. Waisbren, PhD, The Children's Hospital, Gardner 648, 300 Longwood Avenue, Boston, MA 02115. Dr. Shiloh is a psychologist with the Institute for Child Development and with Tel Aviv University, Tel Aviv, Israel; Ms. St. James is a psychologist with the Biochemical Genetics Unit at The Children's Hospital, Boston; Dr. Levy is Senior Associate in Medicine with the Biochemical Genetics Unit at the hospital and also with the Department of Neurology at Harvard Medical School. This paper, submitted to the Journal October 31, 1989, was revised and accepted for publication August 13, 1990.

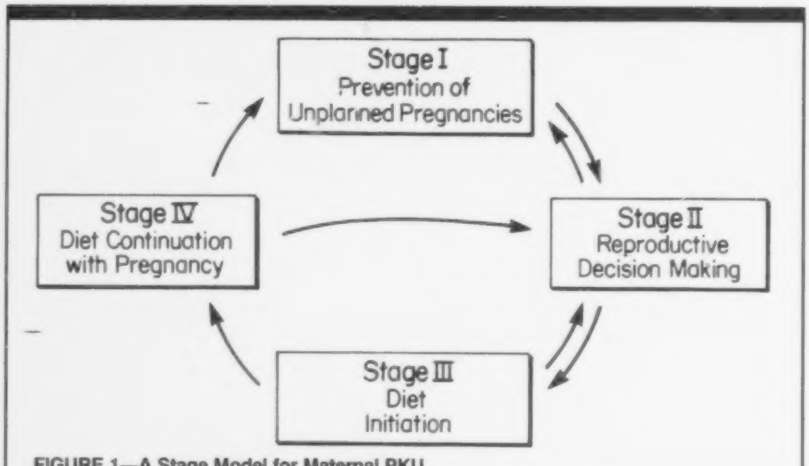
Stage I
Prevention of Unplanned Pregnancies
Stage II
Reproductive Decision Making
Stage III
Diet Initiation
Stage IV
Diet Continuation with Pregnancy

FIGURE 1—A Stage Model for Maternal PKU

Women with moderate or severe mental retardation were excluded from the sample. Recruiting from clinic lists, newborn screening records, and public health registries, we contacted all women with PKU who met the inclusion criteria regarding age and intellectual functioning. We developed two comparison groups—one consisting of age-matched female acquaintances recruited among those whose names were supplied by the subjects and the other consisting of young women with diabetes, who share childbearing risks and the need for medical intervention prior to and during pregnancy with PKU women.[12,13] Each diabetic woman, who had been diagnosed for at least one year and had no major medical complications, was individually matched to a PKU subject for age, and whenever possible, for level of education and marital status.

Two hundred and sixty-two women satisfied the criteria for inclusion in the longitudinal study, and 205 agreed to participate. These included 94 percent of the eligible women with PKU, 74 percent of the acquaintances, and 62 percent of eligible diabetics. Among the 205 who participated, 186 were sexually inactive or sexually active but not desiring pregnancy and were enrolled in the phase of the study concerned with preventing pregnancy. These 186 consisted of 60 with PKU, 65 acquaintances, and 61 with diabetes.

Among the 60 subjects with PKU, 53 have the more severe biochemical defects (classic or atypical PKU) that will require resumption of treatment prior to pregnancy.[3] Seven of the subjects have mild hyperphenylalaninemia, a less severe biochemical defect often left untreated during childhood. Because it is unclear whether mild hyperphenylalaninemia requires dietary treatment during pregnancy,[14,15] these women were included in the PKU group.

Most subjects with PKU had received a written brochure about maternal PKU and had been invited to maternal PKU education groups. Thirty-eight women had attended at least one group. However, the majority had not attended a PKU clinic within the year prior to the study. As a rule, the women had been advised to plan their pregnancies, begin diet, and reduce blood phenylalanine levels before discontinuing contraception. A few had been advised not to have children and to plan on adoption or to wait until more was known about maternal PKU.

Fifty-eight of the 61 women with diabetes (95 percent) are insulin dependent (Type I), while the other three have been treated only with diet (Type II).

Variables and Procedures

Variables for this study were derived from a psychosocial model we developed to describe the challenges faced by women with PKU during various stages in the maternal PKU life-cycle[16] (see figure 1). These variables were selected from a variety of theoretical approaches and covered background issues, personality, knowledge, attitudes and beliefs, and social support. We chose our initial battery after first developing a list of hypotheses regarding prevention of unplanned pregnancy in maternal PKU. This initial battery was piloted extensively among teenagers and women with no known medical conditions, women with diabetes, and older women with PKU.

The dependent (outcome) variables for those women who reported having had sexual intercourse (the sexually active group) were frequency of birth-control use, extent of social activities, and a general quality-of-life score. Frequency of birth-control use refers to the regularity with which a woman reported using any contraceptive method. She indicated this by selecting the frequency of use, on a one to seven scale, ranging from "never" to "all the time." The type of birth control method was related to frequency, with oral contraceptives generally being used "all the time" and all other methods being used less often. Hence, for data analysis purposes, the "frequency" rather than "type" of birth control was selected.

The second outcome variable, extent of social activities, measures the types of activities the woman listed that involve social interactions, such as membership in clubs, participation in religious groups, sports, dating, and parties. This variable was selected to determine if women with PKU avoided certain social situations, perhaps due to anxiety about maternal PKU.

The third outcome variable, the General Health Rating Index (GHRI),[17] measures the degree to which people worry about their health, experience pain, or alter their life-style for health reasons. These topics were also explored in open-ended questions. The responses to the open-ended questions were categorized according to whether the answer was neutral, positive, negative, or a combination of these. All ratings were double-scored, and reliability was greater than 95 percent.

For women who reported never having had sexual intercourse (the sexually inactive group), the social activities scores and the GHRI were the outcome measures.

The independent variables included demographic factors, level of intelligence, knowledge about family planning, knowledge about maternal PKU, personality characteristics, social support for family planning, and attitudes and beliefs about contraception, sex, and childbearing.

The interview consisted of introductory questions that provided demographic information. Socioeconomic status was determined using the Hollingshead-Redlich Two Factor Index of Social Position, based on education and job title.[18] Intelligence was measured by the Wechsler Adult Intelligence Scale-Revised (WAIS-R)[19] or the four-subtest short form of the WAIS-R,[20] when a full-scale score was not available from clinic records.

A test of fertility and contraception knowledge, adapted from Kempton and Forman[21] and a test of knowledge about

TABLE 1—Characteristics of the Subjects

	Study Group	Comparison Groups	
	PKU	Diabetes	Acquaintances
Number	60	61	65
Age (mean ± SD)	20.8 ± 3.6	21.2 ± 3.6	20.8 ± 3.9
Married	7% (4)	12% (7)	11% (7)
Experienced pregnancy	8% (5)	33% (20)	25% (16)
Have child/children	3% (2)	13% (8)	15% (10)
Live with parents	83% (50)	64% (39)	59% (38)
Have driver's license	58% (35)	84% (51)	74% (48)
Education (mean ± SD)	11.9 ± 1.7	13.2 ± 2.1	12.8 ± 1.8
IQ (mean ± SD)	86 ± 14	96 ± 11	96 ± 14

TABLE 2—Sexual Activity and Type and Frequency of Birth Control Use among Women with PKU and Comparison Populations

	Study Group		Comparison Groups			
	PKU		Diabetes		Acquaintances	
	%	(n)	%	(n)	%	(n)
Not Sexually Active	57	(34)	29	(18)	34	(22)
Sexually Active	43	(26)	71	(43)	66	(43)
Method of Birth Control						
Oral Contraceptives	46	(12)	32	(14)	51	(22)
Condom	19	(5)	14	(6)	14	(6)
Diaphragm	5	(1)	9	(4)	5	(2)
Sponge	4	(1)	7	(3)	7	(3)
Tubal Ligation	4	(1)	7	(3)	2	(1)
Rhythm	4	(1)	5	(2)	2	(1)
Intrauterine Device	0		0		0	
No method	19	(5)	26	(11)	19	(8)

maternal PKU[22] (and a comparable test for women with diabetes) were devised for this study. These tests gave one point for a correct response, with 13 possible points for the contraception knowledge test and ten points for the knowledge of maternal PKU and diabetes tests.

Personality was measured by the Adult Nowicki-Strickland Internal-External Locus of Control Scale[23] and the Rosenberg Self-Esteem Scale.[24]

Social support for birth control was determined, as suggested by Fishbein[25] and Vinokur-Kaplan,[26] by asking subjects to name the three people to whom they felt closest and then to rate on a seven-point scale the degree to which each of these people would support their using birth control. The mean ratings constituted the score for social support for birth control.

Attitudes and beliefs were defined by a series of statements to which subjects indicated the extent of their agreement on seven-point scales. These statements, adapted from published studies, were about sexual activity, use of birth control, attitudes regarding contraception, and past pregnancies.[27,28] A composite "attitude towards birth control" score was developed from responses to three questions regarding the benefits, advisability, and comfort in using and procuring contraception.

Each woman in the study was interviewed either in her home or at the hospital. All written questionnaires required no more than a sixth-grade reading level. To increase the comfort of the subjects (and to ensure confidentiality), response cards for personal questions concerning sexual activity, birth control, and pregnancy were provided. Each session lasted about two hours.

Data collection also included review of medical records for information on diagnosis and treatment of PKU or diabetes.

Data were analyzed utilizing the Statistical Package for the Social Sciences (SPSS).[29] To determine differences between the study groups, chi-square or one-way analyses of variance (with Scheffé tests for pairwise comparisons) were used. Pearson correlations were used to identify variables related to successful achievement of outcome goals.

Consent procedures standardized by the Children's Hospital (Boston) Committee on Clinical Investigations were followed.

Results

Characteristics of the Sample

Demographic and other background characteristics of the three groups are listed in Table 1. The groups are similar in age, with a mean of approximately 21 years at the time of the study. The frequencies of those who had been pregnant or had a child were less among the women with PKU. Women with PKU were more likely to live with their parents, were less likely to have a driver's license, had fewer years of education, and had a lower mean IQ. Six women with PKU and one acquaintance (but none with diabetes) performed in the range of mild mental retardation, with scores between 60 and 67.

Prevention of Unplanned Pregnancies

As noted in Table 2, 43 percent of the women in the PKU group reported that they were currently or had been sexually active, a smaller proportion than those in the other groups. Among PKU women who had had sexual relations, 19 percent never used birth control, and 65 percent reported always using birth control when they had sexual relations. These findings, along with the type of contraception selected, were similar to those of the comparison groups. As noted in Table 3, married women with PKU were more likely to always use birth control, while unmarried women were less consistent in this respect.

Women with PKU who reported never having had sexual intercourse and those who were sexually active were similar in age, with mean ages of 20.6 and 20.9 years respectively. Less than 13 percent of both groups indicated disapproval of premarital sex and birth control.

The seven women with mild hyperphenylalaninemia did not differ from other women with PKU in terms of the demographic or psychological variables studied, except that they had a more internal locus of control (F = 9.3, SE = .39). All but two lived with their parents. Their ages ranged from 16 to 22, and their IQ ranged from 78 to 117. Five had had sexual intercourse, among whom two relied on

TABLE 3—Frequency of Birth Control Use among Sexually Active Women according to Marital Status

	Study Group	Comparison Groups	
Use of Birth Control	PKU	Diabetes	Acquaintances
Unmarried Women	22	36	36
Never Uses	5	9	6
Always Uses	13	25	27
Married Women	4	7	7
Never Uses	0	2	2
Always Uses	4	3	5

TABLE 4—Pearson Correlations between Independent Variables and Reported Frequency of Contraceptive Use among Sexually Active Women

	Study Group	Comparison Groups	
Independent Variable	PKU	Diabetes	Acquaintances
Social Support for Birth Control	.64***	.25	.46**
Attitudes about Birth Control	.66***	.22	.16
Knowledge of Family Planning	.43*	.16	.15
Knowledge of Maternal PKU/Diabetes	.03	.17	—
Personality: Locus of Control	–.14	–.39**	.07
Personality: Self-Esteem	–.08	–.17	.08
IQ	.16	.01	.03
Social Class	0	–.12	–.13

NOTE: Results were similar when analyses were redone controlling for age.
*p < .05
**p < .01
***p < .001

TABLE 5—Quality of Life among Sexually Active vs. Inactive Women with PKU

	Sexually Active	Sexually Inactive
General Health Ratings Index (mean ± SD)	82.4 ± 16.5	82.6 ± 15.2
Percent women engaged in spare-time activities		
athletics	34.6	38.2
social activities	57.7	82.4
intellectual pursuits	19.2	23.5
solitary	84.6	79.4
Percent reporting alterations in social life due to PKU	19.2	8.8
Percent reporting alterations in plans for future due to PKU	57.7	58.8
Percent reporting alterations in plans to have children	42.3	55.9

the use of condoms intermittently and three used oral contraception.

Factors Associated with Prevention of Unplanned Pregnancy

As noted in Table 4, the degree of social support for birth control that PKU women received from individuals in their social network and their own attitudes about birth control, including comfort in discussing birth control with a health care provider, were critical factors for contraceptive behavior, correlating .64 and .66 respectively with frequency of contraceptive use. In addition, knowledge of family planning, but not knowledge of maternal PKU, correlated with frequency of contraceptive use. Among women with PKU, personality variables, age, marital status, living situation, IQ, and social class were unrelated to contraceptive use. Multiple regression analyses indicated that when the variable for either social support for birth control or the attitudes about birth control was entered first in the equation, no other variables were significant.

A different pattern of variables predicted contraceptive use for the comparison groups, with only locus of control figuring prominently for the diabetics and only social support for birth control being important for the acquaintances.

Compared with their acquaintances and women with diabetes, women with PKU attained lower scores on the Knowledge of Family Planning Test (F = 7.8 SE = .14) and had a more external locus of control (F = 9.3 SE = .4).

Quality of Life

As noted in Table 5, for women with PKU, the scores on the GHRI were comparable for those who were sexually active and those who were not. In general, they were involved in similar types of spare-time pursuits, although the women who reported never having had sexual relations were somewhat more likely to engage in social activities. Few of the sexually active or sexually inactive women reported alterations in their social lives because of PKU. Women with PKU who identified negative effects from their disease reported having fewer friends, fewer activities, difficulties explaining their disease to boyfriends, and concerns about becoming sexually involved. More than half in each group reported alterations in their plans for the future due to PKU. Somewhat more of the sexually inactive women mentioned concerns about maternal PKU, and equal numbers mentioned reduced career opportunities as a consequence of PKU-related learning disabilities. Nonetheless, compared with women with diabetes, women with PKU attained significantly higher scores on the GHRI (F = 16.5 SE = 1.5), indicating less worry, less pain, and less of a negative impact from their disease. Fifty percent of women with PKU, compared with 85 percent of women with diabetes, reported undesirable alterations in their social activities and plans for the future as a consequence of their disease.

Discussion

Until now studies of individuals with PKU have focused on intellectual and academic development. However, the necessity for treatment prior to pregnancy in maternal PKU stimulated us to investigate other aspects of development. In this study, we concentrated on those psychosocial factors in young women with PKU that we believe will be important in their meeting the challenges of family planning in maternal PKU.

We found that many young women with PKU tend to live with their parents and to be without a driver's license. Nonetheless, they are engaged in social activities and competitive employment. They plan to marry and bear children. It is unclear why fewer women with PKU reported having had sexual intercourse compared with their acquaintances and the diabetic women, who were similar in age and marital status. We believe, however, that over the next few years, more will become sexually active. Unfortunately, those who are sexually active seem to be at high risk for unplanned pregnancies, since 35 percent of them report only intermittent use of birth control.

It appears that social support for birth control and positive attitudes about contraception each have a very significant influence on the reported contraceptive behavior of young women with PKU. These findings are consistent with Fishbein's theory of reasoned action, which states that attitudes, social norms, and the willingness to comply with these norms are predictors of behavior.[30-32]

With the exception of social support, those factors that correlate with birth control use in women with PKU are specific to them and unlike correlating factors in the two comparison groups studied. For instance, their external locus of control does not predict contraceptive frequency; yet in the diabetic subjects and in published studies, an external locus of control is associated with infrequent contraceptive use.[33-34] On the other hand, knowledge of family planning is significantly correlated with reported frequency of birth control use only in the PKU women. Their knowledge of family planning, in general, is lower than that of the other subjects studied and, in many instances, may be sufficiently low to impair their ability to choose and use contraceptive measures. Education programs such as those described by Cohen and his colleagues in Israel[35] may be beneficial to these women. In developing our test of knowledge of maternal PKU, we found that knowledge scores were higher among the 38 women who had attended a PKU education group than among the 22 who had not done so.[22] Finally, although the women with PKU have attitudes about birth control that are comparable to those of their acquaintances and the diabetic women, their contraceptive behavior appears to be more highly influenced by their attitudes. Further analyses show that although IQ is significantly related to knowledge of family planning and knowledge of maternal PKU, it is not significantly related to the frequency of contraceptive use.

In addition to these findings, we gained a keener sense of how women with PKU approach the question of childbearing. The vast majority consider having children a natural sequel to finding a husband, and few acknowledge that an unplanned pregnancy might occur. In these attitudes, they are similar to adolescents confronting a variety of medical and health issues.[36]

We also learned more about the ways in which women with PKU related to health care providers. As a group, they were extraordinarily cooperative, perhaps reflecting a sense of trust developed over years of association with the PKU programs. Ninety-four percent of those contacted participated in the study. Despite this cooperation, only a few of the women made follow-up clinic appointments. This may be due to apprehension about the blood test and physical examination, to a belief that they already know enough about maternal PKU, or to a presumption that an unplanned pregnancy was not likely to occur.

Women with PKU were self-accepting, as indicated by their responses to the self-esteem questionnaire. Many expressed the thought that they "used to have" PKU when they were children on the special diet. For them, prevention of unplanned pregnancies may have been no different than it was for other teenagers and young adults striving for a sense of identity and independence. The risks to the fetus could easily be forgotten or ignored during this time of self-exploration. This may explain why birth control practices among those with PKU were no better than those of the acquaintances who did not face known childbearing risks. On the other hand, fewer women with PKU were sexually active, perhaps due to fears of unplanned pregnancies, to concerns regarding risks to the fetus, or to limited social opportunities. These speculations regarding the psychological impact of PKU will be explored further as the study continues. Social support will also receive more intensive examination in the continuing study. We plan to define social support more precisely in terms of the nature of support networks, the importance of peers versus family, the role of perceived as opposed to actual support provided, and the extent to which the concept reflects general support as compared with specific encouragement for birth control use.

As in any study dependent on self-report, the reliability and validity of the subjects' responses should be checked. We asked several duplicate questions, worded slightly differently, so that inconsistent answers about sexual activity and birth control use could be verified. Our sense was that the women were generally honest. Since this is a longitudinal study, each subject will be interviewed at least five times, and further checks can be made on the consistency and reliability of the results.

In the near future as more girls with PKU enter childbearing ages, the need for specific programs for maternal PKU, already serious, will dramatically increase. A critical first step is to develop a registry in the United States similar to that in Canada to track women with PKU so that information and services can be provided. Our society and its health care system must confront the issue of maternal PKU. To ignore the issue is to punish those who because of their inherited condition are at risk for bearing defective children. ☐

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Conference of Public Health Laboratorians 1991 Kimble Methodology Award

The Kimble Methodology Award, sponsored by the Conference of Public Health Laboratorians and supported by Kimble Glass, Inc., Vineland, New Jersey is intended to recognize significant contributions to the development of diagnostic methodology for use in the public health or diagnostic laboratory. The 1990 Award winners were Dr. Robert E. Weaver and Ms. Dannie G. Hollis, Special Bacteriology Laboratory, Centers for Disease Control, Atlanta, Georgia, for their work on the characterization and identification of numerous genera and species of medically important bacteria.

Nominations are being sought for the 1991 Kimble Methodology Award, (the 40th annual award) to be presented at the annual meeting of the Conference of Public Health Laboratorians scheduled for November 10, 1991 in Atlanta, Georgia. The award consists of a $1,000 check and a plaque presented by Kimble Glass. In addition, travel and subsistence will be paid for the recipient to attend the 1991 Conference meeting.

The deadline for submission of nominations for the 1991 award is June 17, 1991. All nominations should be entered according to the following rules:

- Nominee's work should be a fundamental contribution, or an adaptation of such, to diagnostic methodology for use in the public health or diagnostic laboratory. Degree of acceptance and use of the methodology will be considered.
- Nominee(s), either an individual or two or more working together, must live and work on the North American continent.
- The nomination (10 copies) shall consist of a letter describing the reason for nomination; the recognition accorded the nominee for the work and the degree of acceptance by and/or its importance to the work of public health or diagnostic laboratories; the curriculum vitae of the nominee; the nominee's current place of employment; a list of the nominee's publications along with reprints of representative recent publications; and, optionally, letters of support for the nomination.

Send nominations to: Hilton T. Goulson, PhD, Kimble Award Chairperson, Department of Parasitology and Laboratory Practice, UNC School of Public Health, CB #7400, Chapel Hill, NC 27599-7400. Tel: (919) 966-3892.

Predicting Onset and Chronicity of Women's Problem Drinking: A Five-Year Longitudinal Analysis

ABSTRACT

Background. Longitudinal studies of adult drinking have typically excluded or sampled only small numbers of problem drinking women, and have measured a limited range of influences on women's drinking behavior.

Method. To study the development of women's problem drinking over time, five-year follow-up interviews were conducted with two groups of respondents from a 1981 national survey of women's drinking: 143 problem drinkers and 157 nonproblem drinkers. Regression analyses examined effects of 1981 predictors on six measures of 1986 problem drinking, for problem drinkers and nonproblem drinkers separately.

Results. Among 1981 nonproblem drinkers, predictors of onset of problem drinking indicators by 1986 included younger age, cohabiting, and lifetime use of drugs other than alcohol. The most consistent predictor of persistent (chronic) problem drinking was sexual dysfunction; other predictors included being employed part-time or never married, and experiencing recent depression. Divorce or separation predicted *lower* levels of subsequent alcohol dependence among problem drinkers.

Conclusions. Findings suggest that different personal and social factors predict the onset of problem drinking as compared with its continuation, and point to nontraditional life-style, sexual dysfunction, and role deprivation as potentially important variables. (*Am J Public Health* 1991;81:305–318)

Sharon C. Wilsnack, PhD, Albert D. Klassen, MA, Brett E. Schur, PhD, and Richard W. Wilsnack, PhD

Introduction

Longitudinal research on general population samples is essential for predicting the onset and chronicity of problem drinking, but has not been adequately applied to the study of women drinkers. A majority of the longitudinal studies in a recent comprehensive review[1] totally excluded female respondents. Of 12 longitudinal studies of adult drinking included in an international collaborative effort,* four have exclusively male samples, while only one focuses primarily on females.

Those longitudinal studies that do sample women frequently have insufficient numbers of heavy drinking and problem drinking women for multivariate analysis, due to the lower rates of heavy drinking and drinking-related problems among women than among men. Even a relatively large national probability sample of 2,000 respondents (half female) yields only 30 to 60 women who drink as heavily as 14 drinks per week.[2] One frequently cited longitudinal study[3] had only three women in its problem drinker sample. In addition, most longitudinal studies that include women have focused on immediate contexts and correlates of drinking behavior, with relatively little attention to other or longer term influences on women's drinking.

The five-year longitudinal study reported here gathered information about diverse antecedents and consequences of drinking and drinking problems in women. This paper addresses two major questions: 1) What factors predict the incidence (onset) of problem drinking among women over a five-year period? 2) What factors predict the chronicity—continuation or worsening—of women's problem drinking over five years? Because drinking may be reciprocally related to a number of personal and social-environmental conditions, we analyze drinking as both a predictor and an outcome of these conditions.

Methods

Sampling

The data are from a national survey of women's drinking conducted in the fall and winter of 1981 and from a follow-up survey in the fall and winter of 1986. Sampling and fieldwork for both surveys were conducted by the National Opinion Research Center (NORC).

The 1981 survey interviewed a stratified probability sample of the noninstitutionalized US population age 21 and over, obtained from initial screening of 4,032 households. After 10-minute screening interviews to determine women's drinking levels, NORC interviewers requested interviews from every moderate-to-heavy drinking woman (four or more drinks per

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Address reprint requests to Sharon C. Wilsnack, PhD, Professor, Department of Neuroscience, University of North Dakota School of Medicine, 501 N. Columbia Road, Grand Forks, ND 58203. Mr. Klassen and Dr. R. Wilsnack are also with the Department of Neuroscience at the School of Medicine; Dr. Schur is with the Crozer-Chester Medical Center, Upland, Pa. This paper, submitted to the Journal December 8, 1989, was revised and accepted for publication August 6, 1990.

week), from every woman reporting a history of drinking-related problems, and from one of every four lighter drinking or abstaining women. Interviews were completed with 500 moderate-to-heavy drinking women and 39 women who were self-reported former problem drinkers (89 percent completion rate), and with 378 lighter drinking and abstaining women (83 percent completion rate). A probability sample of 396 men (66 percent completion rate) was also interviewed as a comparison group. Detailed information about sample design, statistical weighting, and comparisons with other national data are provided elsewhere.[2,4]

The 1986 follow-up survey compared two subsamples of female respondents. Problem Drinkers (PDs) were women who in 1981 had reported at least two of the following: 1) average consumption of one or more ounces of ethanol per day; 2) one or more drinking-related problems in the past 12 months; and 3) one or more alcohol dependence symptoms in the past 12 months. These problem drinking criteria are similar to those used in other studies,[5-7] and the proportion of women thus identified—10.1 percent of the total weighted female sample—falls in the middle of the range of estimates of problem drinking rates among women in US surveys during the 1960s and 1970s (1 percent to 21 percent).[6] Nonproblem Drinkers (NPDs) were women who in 1981 drank more than one drink per month but reported none of the three problem drinking indicators.

Application of these criteria produced samples of 178 PDs and 199 NPDs. Of these, 157 PDs (88.2 percent) and 185 NPDs (93.0 percent) were successfully located by NORC staff. Six percent of the total follow-up sample (five PDs and 16 NPDs) were deceased, too ill or senile to be interviewed, or out of the country. Of the remaining respondents potentially available for interviews, interviews were completed with 143 PDs (94.1 percent) and 157 NPDs (92.9 percent). The final sample thus represents 80.3 percent (143 of 178) of all PD cases in the 1981 sample and 78.9 percent (157 of 199) of all 1981 NPDs. (Not included in the follow-up sample were 290 women who abstained at least 30 days before the 1981 survey, 110 infrequent drinkers who drank one drink per month or less in the 12 months preceding the survey, and 140 women who drank more than one drink per month and who reported only one of the three problem drinking indicators.)

In order to assess possible selection bias due to attrition, nonresponse rates (combining refusals and locating failures) were examined for respondent categories based on all major drinking variables, seven major demographic variables, and two sexuality variables (nontraditional sexual behavior and sexual dysfunction), within PD and NPD subsamples separately. Nonresponse rates for the following categories exceeded the overall subsample nonresponse rates by 15 percent or more: Black PDs; NPDs age 71 and older; PDs separated from their husbands; PDs with household income under $8,000; NPDs with household income $8,000–$15,000; NPDs completing college but not graduate school; NPDs who were retired or disabled; and NPDs reporting both masturbation and premarital sexual experience. None of the drinking variables had categories with excessive nonresponse rates. Based on their known or expected relationships to other variables, income and race seemed to be the two most important variables for weighting adjustments. Accordingly, nonresponse weighting adjustments were applied to the 1986 sample based on division of the PD and NPD subsamples into three categories: Blacks; race Other/income under $15,000; race Other/income $15,000 and above.

Data Collection

The 1981 survey questionnaire was administered face-to-face by female NORC interviewers. It included questions (described below) about drinking behavior, drinking-related problems, alcohol dependence symptoms, and lifetime drinking history. Other questions dealt with demographic characteristics, family history and socialization, personality traits and values, social roles, characteristics of the respondent's primary interpersonal relationship, stressful life events, depression, sexual experience, reproductive dysfunction, use of drugs other than alcohol, and antisocial behavior.

The follow-up survey, again administered by female NORC interviewers, used measures of drinking behavior and drinking consequences identical to those used in 1981. Other questions asked about changes in demographic characteristics such as marital and employment status since 1981, and about other major life events such as births, deaths, physical illness, depression, and sexual and reproductive dysfunction. New questions addressed employment experiences, feelings of role overload and role deprivation, sexual experience and sexual abuse, and relationship conflict/violence.

Measures

Predictor Variables: The 1981 survey was designed to assess the individual and combined effects of a large number of variables that had been associated with women's drinking in previous theory and research, as well as additional variables not addressed in earlier work. The conceptual model guiding the study assumed that a woman's demographic and family background (including such variables as age, education, religious orientation, and recalled parental behavior including drinking) affects three major sets of influences on drinking behavior: 1) personality characteristics (including personal motives and values, drinking-related beliefs and attitudes, and gender-role attributes); 2) perceived social environment (including social norms and interactions with significant others); and 3) life experiences (including stressful life events, sexual and reproductive experience, affective disorder, other drug use, and antisocial behavior). Background, personality, social environment, and life experiences all influence contexts of drinking behavior, levels of alcohol consumption, and drinking consequences.[8] From the large set of variables initially included in the 1981 survey, variables were selected for the present longitudinal analyses if they had shown a significant positive relationship to heavier drinking and/or to adverse consequences of drinking in earlier cross-sectional analyses of 1981 data.[2,4,8-13]

Time 1 predictors were demographic characteristics, personality attributes, social-environmental conditions, and life-history events as measured in the 1981 survey. For some variables, additional interim predictors consisted of specific changes in demographic status (e.g., becoming divorced or unemployed) or major life experiences (e.g., miscarriage/stillbirth, depressive episode) during the 1981–86 follow-up interval. We also measured the Time 2 (1986) values of Time 1 predictors in order to analyze possible Time 2 consequences of Time 1 drinking. Brief descriptions of personality variables, sexuality-related variables, and drinking variables are provided in the Appendix; descriptions of all other variables can be obtained by writing the authors.

Drinking Variables: Questions about the frequency and quantity of consumption of wine, beer, and liquor were used to estimate respondents' average daily intake in ounces of ethanol per day for the 30 days preceding the survey.[5,14] Minor refinements in the 1986 survey adjusted for

TABLE 1—Changes in Problem Drinking Indicators, 1981–86, among 1981 Nonproblem Drinkers (NPDs) and Problem Drinkers (PDs), by Age (%)

1986 Problem Drinking Indicators	Age 21–34		Age 35–49		Age 50+		Total	
	NPDs	PDs	NPDs	PDs	NPDs	PDs	NPDs	PDs
None	73.0	37.8	95.6	25.8	95.4	26.1	88.8	32.7
One	13.8	34.8	4.4	17.1	3.4	38.3	6.9	30.7
Two or Three	13.3	27.4	0	57.2	1.2	35.6	4.3	36.5
Weighted N	(46)	(80)	(58)	(37)	(49)	(23)	(153)	(140)

the new popularity of wine coolers and low-alcohol beer, the use of malt liquor, and slight declines in the ethanol content of wine and liquor. Estimating procedures are summarized in the Appendix. Other drinking measures included the frequency of heavy episodic drinking (six or more drinks per day) and the frequency of intoxication ("feeling drunk") in the 12 months preceding the survey.

An index of nine kinds of problem consequences of drinking included driving while feeling intoxicated, starting fights (with family members or others), damage to job performance, interference with housework or chores, drinking-related home accidents and problems with children, and spouse's or partner's complaints about the respondent's drinking and threats to leave the drinker. An index of five alcohol dependence symptoms consisted of memory lapses while drinking (blackouts), rapid drinking, morning drinking, inability to stop drinking before becoming intoxicated, and inability to stop or reduce alcohol consumption over time. These indexes summed the number of different consequences and symptoms reported for the 12 months preceding the survey. A composite Problem Drinking Index (PDI) summed the occurrence in the past 12 months of 1) any episode of intoxication, 2) any problem consequence, and 3) any alcohol dependence symptom (scores ranged from zero to three).

Data Analysis

Data analysis involved weighting cases to take into account unequal probabilities of inclusion at several levels of selection, including 1981 stratification by drinking level, 1986 selection of PD and NPD samples, and response rate variations in both surveys.[4] For significance tests we adjusted weights by constant ratios to produce weighted Ns equal to the actual numbers of respondents in the PD and NPD samples.

This paper reports results of three analyses. First, cross-tabulations describe patterns of change in problem drinking indicators between the 1981 and 1986 surveys. Second, regression analyses examine the effects of individual Time 1 and interim predictors on six Time 2 drinking variables, and the effects of six Time 1 drinking variables on Time 2 values of Time 1 predictors. Dependent variables in the regression analyses were standardized residuals. Each Time 2 drinking variable was regressed on its Time 1 equivalent plus measures of 1981 consumption (average consumption and heavy episodic drinking) to produce standardized residuals as measures of Time 2 drinking patterns not predicted by Time 1 drinking patterns. Similarly, each Time 2 condition (e.g., sexual dysfunction in 1986) was regressed on its Time 1 equivalent, producing standardized residuals as measures of Time 2 conditions not accounted for by Time 1 conditions.

The third analysis uses results of the second analysis as a winnowing procedure to develop 12 unique multiple regression equations showing the combined effects of Time 1 and interim predictors on the same standardized residuals of each of six measures of Time 2 drinking, for the NPD and PD samples separately. Predictors were included in these equations 1) if they predicted ($p < .10$) the Time 2 drinking variable in the bivariate analyses described above, or 2) if they predicted ($p < .10$) the Time 2 drinking variable in (a) equations that entered all Time 1 predictors simultaneously, or (b) equations that entered all predictors simultaneously, using interim predictors for those variables with both Time 1 and interim versions. Every equation included age as a control variable. In the 12 final equations, predictors were entered hierarchically in two steps: Time 1 predictors (those that occurred before 1981 or that constituted the respondent's status in 1981), followed by interim predictors (occurring during the follow-up interval). Based on earlier cross-sectional findings, every predictor (except age) in the second and third analyses was hypothesized to predict heavier drinking or more adverse drinking consequences. For this reason, one-tailed tests of significance were used for relationships in the predicted direction, two-tailed tests for relationships in the opposite direction.

Design effects were assessed for seven of the resulting regression equations.[15] In only two (intoxication and PDI, among PDs only) did the mean of the individual square roots of design effects slightly exceed 1.0 (1.017 and 1.101 respectively), and the overall mean for all predictors in the seven equations was less than 1.0 (0.963). This indicates that the findings reported here are not affected by weighting or cluster characteristics of the PD and NPD subsamples to an extent that would require adjustment for departures from the simple random sampling model.

Results

Changes in Problem Drinking

Table 1 shows five-year changes in the three problem drinking indicators (consumption, problems, symptoms) used to define the 1981 NPD and PD samples. A large majority of all 1981 NPDs remained free of problems five years later; 11.2 percent had developed at least one problem drinking indicator. Of the 1981 PDs (who in 1981 had reported at least two of the three indicators), approximately one-third (32.7 percent) by 1986 reported none of the three indicators and another one-third (30.7 percent) reported only one indicator.

Patterns of change were quite different for three age groups:

- Onset of problem drinking was most common among young women age 21–34; more than one-fourth (27.1 percent) of the 1981 NPDs in this age group reported at least one problem drinking indicator by 1986, compared with 4.4 percent of 1981 NPDs age 35–49 ($\chi^2 = 8.85$,

TABLE 2—1981 Sociodemographic Predictors of Five-Year Onset and Chronicity of Problem Drinking Indicators: Hierarchical Multiple Regression Coefficients with Unadjusted (Zero-Order) Weighted Means and Rates

	Time 2 (1986) Problem Drinking Indicators						
Time 1 (1981) Sociodemographic Predictors	Average Consumption (ounces ethanol/day)	Heavy Episodic Drinking (% any)	Intoxication (% any)	Problem Consequences (% any)	Dependence Symptoms (% any)	Problem Drinking Index (% 1⁺)	Problem Drinker Sample Criteria (%1⁺)
Younger Age							
NPD Onset							
Beta 1	−.04	+.22	+.30	+.17	+.10	+.23	*
Beta 2	−.08	+.20	+.25	+.06	+.04	+.13	
21–34 (51)	.175	35%	52%	24%	14%	54%	27%
35+ (106)	.151	10%	6%	3%	1%	7%	4%
PD Chronicity							
Beta 1	−.05	+.15	+.15	+.02	+.05	+.11	
Beta 2	−.21	+.13	+.15	+.01	+.07	+.12	
21–34 (77)	.490	57%	75%	53%	37%	78%	62%
35+ (66)	.881	59%	63%	54%	44%	76%	74%
Unemployed, 1981							
NPD Onset							
Beta 1	—	—	—	+.07	—	—	
Beta 2	—	—	—	−.01	—	—	
Yes (10)	.213	37%	28%	28%	11%	28%	28%
No (146)	.151	17%	19%	9%	5%	21%	11%
PD Chronicity							
Beta 1	—	—	—	—	—	—	
Beta 2	—	—	—	—	—	—	
Yes (7)	.457	56%	73%	44%	49%	73%	62%
No (135)	.668	58%	70%	54%	39%	77%	67%
Employed Part Time, 1981							
NPD Onset							
Beta 1	—	−.16	—	—	—	—	
Beta 2	—	−.17	—	—	—	—	
Yes (31)	.216	14%	15%	8%	4%	18%	12%
No (125)	.146	18%	20%	10%	5%	22%	11%
PD Chronicity							
Beta 1	+.13	—	—	−.16	—	—	
Beta 2	+.23	—	—	−.14	—	—	
Yes (22)	.674	53%	64%	44%	33%	70%	50%
No (119)	.637	60%	73%	56%	42%	80%	70%
Nontraditional Job for Women, 1981							
NPD Onset							
Beta 1	—	+.07	+.12	—	—	—	
Beta 2	—	+.04	+.06	—	—	—	
Yes (42)	.149	24%	26%	6%	3%	27%	9%
No (107)	.165	15%	15%	8%	4%	17%	9%
PD Chronicity							
Beta 1	—	—	−.17	—	—	—	
Beta 2	—	—	−.14	—	—	—	
Yes (47)	.630	64%	64%	52%	38%	71%	72%
No (80)	.613	55%	78%	53%	42%	84%	61%
Never Married, 1981							
NPD Onset							
Beta 1	+.06	−.01	—	—	+.12	—	
Beta 2	+.08	+.04	—	—	+.12	—	
Yes (17)	.242	34%	34%	18%	21%	34%	26%
No (140)	.151	16%	18%	9%	4%	20%	10%
PD Chronicity							
Beta 1	—	+.05	+.18	+.17	—	+.18	
Beta 2	—	+.05	+.16	+.16	—	+.14	
Yes (22)	.519	74%	86%	76%	52%	92%	87%
No (121)	.686	55%	67%	49%	38%	74%	64%
Divorced/Separated, 1981							
NPD Onset							
Beta 1	—	−.14	—	—	—	—	
Beta 2	—	−.14	—	—	—	—	
Yes (18)	.135	10%	34%	9%	3%	35%	9%
No (139)	.162	19%	17%	10%	5%	19%	12%

(continued)

TABLE 2—Continued

Time 1 (1981) Sociodemographic Predictors	Time 2 (1986) Problem Drinking Indicators: Average Consumption (ounces ethanol/day)	Heavy Episodic Drinking (% any)	Intoxication (% any)	Problem Consequences (% any)	Dependence Symptoms (% any)	Problem Drinking Index (% 1^+)	Problem Drinker Sample Criteria (%1^+)
PD Chronicity							
Beta 1	—	—	−.09	—	—	−.08	
Beta 2	—	—	−.08	—	—	−.07	
Yes (29)	.473	45%	55%	48%	23%	66%	57%
No (115)	.709	61%	74%	54%	44%	80%	70%
Cohabiting, 1981							
NPD Onset							
Beta 1	—	+.18	+.08	—	—	—	
Beta 2	—	+.17	+.06	—	—	—	
Yes (5)	.070	61%	51%	13%	0%	64%	13%
No (152)	.161	16%	18%	9%	5%	20%	11%
PD Chronicity							
Beta 1	—	−.26	—	—	—	—	
Beta 2	—	−.22	—	—	—	—	
Yes (14)	.865	24%	76%	52%	30%	76%	52%
No (129)	.633	62%	69%	54%	41%	78%	69%
Recent Empty Nest, 1981							
NPD Onset							
Beta 1	—	—	—	—	+.12	—	
Beta 2	—	—	—	—	+.10	—	
Yes (17)	.164	25%	25%	10%	12%	25%	21%
No (139)	.158	17%	19%	9%	5%	21%	10%
PD Chronicity							
Beta 1	—	—	—	—	—	—	
Beta 2	—	—	—	—	—	—	
Yes (16)	.925	51%	68%	64%	36%	84%	81%
No (124)	.628	58%	70%	51%	40%	76%	65%

*Not analyzed as a dependent variable.

Note.–Time 2 drinking variables were standardized residuals from regressing 1986 drinking variables on their 1981 equivalents. Each column shows multiple regression results for predictors of onset and chronicity of a specific Time 2 drinking variable. Predictors from the same equation appear in Tables 2 (sociodemographic), 3 (personality/life experience), and 4 (interim predictors). Beta 1 coefficients are from the first hierarchical step (including only Time 1 predictors) and Beta 2 coefficients are from the second hierarchical step (adding interim predictors). Underlined betas are significant at $p<.05$. With the exception of the far right-hand column (Problem Drinker Sample Criteria were not analyzed as a dependent variable), the absence of beta values indicates that the predictor was not significantly associated with that drinking variable in preliminary bivariate and multivariate winnowing analyses, and thus was not included in the final hierarchical equation. Weighted means and percentages are provided to illustrate the fit, or contrast, of simple bivariate comparisons with the more rigorous and conservative multiple regression analyses. Although means and rates are weighted, Ns are unweighted to show actual numbers of women in predictor categories. R^2 and R^2 change values for each equation and each hierarchical step are given at the bottom of each column, at the end of Table 4.

df = 1, p = .003) and 4.6 percent of 1981 NPDs age 50 or older ($\chi^2 = 7.54$, df = 1, p = .006).

- Remission of problem drinking was also most common among the youngest women; by 1986 only 27.4 percent of the 1981 PDs in the youngest age group still reported two or three problem drinking indicators, compared with 48.9 percent of 1981 PDs age 35 and older ($\chi^2 = 5.93$, df = 1, p = .015).
- Chronicity of problem drinking, on the other hand, was most likely among women age 35–49. Of the 1981 PDs in this age group, 57.2 percent continued to experience at least two problem drinking indicators in 1986, compared with 27.4 percent of women age 21–34 ($\chi^2 = 8.40$, df = 1, p = .004) and 35.6 percent of women age 50 or older ($\chi^2 = 1.84$, df = 1, p = .18; for 35 to 49-year-olds vs. all others, $\chi^2 = 7.97$, df = 1, p = .005).

Bivariate Relationships between Time 1 Predictors and Time 2 Drinking, and between Time 1 Drinking and Time 2 Conditions

Bivariate analyses consisted of a series of regression equations that used each Time 1 and interim predictor to predict the standardized residuals of each Time 2 drinking variable, and reciprocal analyses that used each Time 1 drinking variable to predict the standardized residuals of Time 2 or interim conditions corresponding to Time 1 predictors. These bivariate analyses showed many significant time-ordered relationships. Of 23 Time 1 variables, ten were significant predictors of Time 2 drinking variables among NPDs, and ten were significant predictors among PDs. Of ten interim variables, five were significant predictors of Time 2 drinking variables among NPDs, and six were significant predictors among PDs. In the other direction, of 19 Time 2 or interim conditions corresponding to Time 1 predictors, eight were significantly predicted by Time 1 drinking variables among NPDs, and 12 were significantly predicted by Time 1 drinking variables among PDs. Detailed results from the bivariate analyses are not presented here because of space limitations but are available upon request from the authors. Tables 2–4 include all Time 1and interim variables that significantly predicted any of the six Time 2 drinking variables among either PDs or NPDs in the bivariate analyses. Of the 33 potential

TABLE 3—1981 Personality and Life Experience Predictors of Five-Year Onset and Chronicity of Problem Drinking Indicators: Hierarchical Multiple Regression Coefficients with Unadjusted (Zero-Order) Weighted Means and Rates

Time 1 (1981) Personality/Life Experience Predictors	Time 2 (1986) Problem Drinking Indicators						
	Average Consumption (ounces ethanol/day)	Heavy Episodic Drinking (% any)	Intoxication (% any)	Problem Consequences (% any)	Dependence Symptoms (% any)	Problem Drinking Index (% 1^+)	Problem Drinker Sample Criteria (% 1^+)
Depressive Episodes up to 1981							
NPD Onset							
Beta 1	—	—	—	—	—	—	
Beta 2	—	—	—	—	—	—	
Any (46)	.147	17%	20%	7%	3%	20%	11%
None (111)	.163	18%	19%	10%	6%	22%	12%
PD Chronicity							
Beta 1	—	—	+.11	+.17	+.04	+.10	
Beta 2	—	—	+.05	+.13	−.01	+.06	
Any (77)	.603	55%	74%	55%	45%	80%	65%
None (66)	.724	61%	66%	51%	34%	74%	70%
Anxiety, 1981							
NPD Onset							
Beta 1	—	—	—	—	—	—	
Beta 2	—	—	—	—	—	—	
Yes (39)	.113	18%	27%	9%	5%	26%	9%
No (118)	.179	17%	16%	10%	5%	19%	12%
PD Chronicity							
Beta 1	−.20	—	—	−.16	—	—	
Beta 2	−.23	—	—	−.17	—	—	
Yes (55)	.565	57%	77%	48%	48%	84%	63%
No (88)	.718	58%	66%	56%	35%	73%	70%
Low Self-Esteem, 1981							
NPD Onset							
Beta 1	—	+.18	+.10	—	—	+.02	
Beta 2	—	+.16	+.10	—	—	+.01	
Yes (47)	.140	28%	26%	10%	3%	25%	11%
No (107)	.162	14%	18%	10%	6%	20%	11%
PD Chronicity							
Beta 1	—	—	—	—	—	—	
Beta 2	—	—	—	—	—	—	
Yes (54)	.622	58%	68%	50%	29%	72%	63%
No (88)	.667	57%	71%	55%	46%	80%	70%
Nontraditional Gender Traits, 1981							
NPD Onset							
Beta 1	—	+.08	—	—	—	—	
Beta 2	—	+.06	—	—	—	—	
Yes (48)	.182	22%	21%	11%	4%	24%	14%
No (106)	.155	17%	20%	9%	6%	21%	10%
PD Chronicity							
Beta 1	+.26	+.12	+.14	+.14	—	+.13	
Beta 2	+.25	+.09	+.15	+.13	—	+.14	
Yes (47)	1.021	58%	75%	60%	44%	82%	77%
No (91)	.464	58%	67%	49%	36%	75%	62%
No. of Drugs Used up to 1981							
NPD Onset							
Beta 1	—	+.13	+.12	+.10	+.18	+.16	
Beta 2	—	+.10	+.14	+.08	+.14	+.15	
2^+ (26)	.196	23%	30%	10%	7%	30%	17%
0–1 (129)	.153	16%	18%	10%	5%	20%	10%
PD Chronicity[a]							
Beta 1							
Beta 2							
2^+ (45)	.875	59%	67%	57%	49%	74%	73%
0–1 (98)	.559	57%	71%	51%	36%	79%	64%
No. of Drugs Used in 1961							
NPD Onset[a]							
Beta 1							
Beta 2							
2^+ (4)[b]							
0–1 (144)	.165	17%	20%	10%	5%	22%	11%

(*continued*)

TABLE 3—Continued

	Time 2 (1986) Problem Drinking Indicators						
Time 1 (1981) Personality/Life Experience Predictors	Average Consumption (ounces ethanol/day)	Heavy Episodic Drinking (% any)	Intoxication (% any)	Problem Consequences (% any)	Dependence Symptoms (% any)	Problem Drinking Index (% 1⁺)	Problem Drinker Sample Criteria (%1⁺)
PD Chronicity							
Beta 1	+.16	—	—	—	+.14	+.11	
Beta 2	+.07	—	—	—	+.14	+.10	
2^+ (15)	1.543	72%	81%	62%	81%	81%	91%
0–1 (125)	.540	55%	68%	53%	35%	76%	65%
Premature Baby, up to 1981							
NPD Onset							
Beta 1	—	—	—	—	—	—	
Beta 2	—	—	—	—	—	—	
Yes (10)	.117	0%	5%	0%	0%	5%	0%
No (145)	.162	19%	20%	10%	5%	22%	12%
PD Chronicity							
Beta 1	—	—	—	+.09	+.06	+.11	
Beta 2	—	—	—	+.08	+.03	+.10	
Yes (13)	.584	74%	82%	52%	54%	87%	73%
No (129)	.670	56%	69%	53%	38%	76%	67%
Miscarriage/Stillbirth, 1976–81							
NPD Onset							
Beta 1	—	+.05	—	—	—	—	
Beta 2	—	+.04	—	—	—	—	
Yes (6)	.082	62%	45%	0%	15%	45%	15%
No (149)	.161	16%	19%	10%	5%	20%	11%
PD Chronicity							
Beta 1	—	—	—	+.01	—	—	
Beta 2	—	—	—	+.03	—	—	
Yes (5)	.265	59%	80%	39%	39%	80%	39%
No (137)	.675	58%	70%	53%	40%	77%	68%
Infertility, up to 1981							
NPD Onset							
Beta 1	—	—	+.10	—	—	—	
Beta 2	—	—	+.11	—	—	—	
Yes (31)	.158	11%	18%	11%	8%	22%	14%
No (125)	.159	19%	20%	9%	4%	20%	10%
PD Chronicity							
Beta 1	—	—	—	—	—	—	
Beta 2	—	—	—	—	—	—	
Yes (18)	.916	78%	73%	49%	46%	73%	76%
No (125)	.625	55%	70%	54%	39%	78%	66%
Liberal Sexual Morality, 1981							
NPD Onset							
Beta 1	—	—	+.02	+.07	—	+.09	
Beta 2	—	—	+.02	+.14	—	+.11	
Yes (71)	.183	24%	35%	19%	10%	40%	22%
No (81)	.139	14%	10%	4%	2%	10%	5%
PD Chronicity							
Beta 1	—	—	—	+.10	—	—	
Beta 2	—	—	—	+.10	—	—	
Yes (95)	.755	63%	72%	58%	40%	79%	71%
No (47)	.475	43%	65%	45%	41%	74%	61%
Nontraditional Sexual Behavior, up to 1981							
NPD Onset							
Beta 1	+.11	+.13	+.10	—	—	+.08	
Beta 2	+.14	+.15	+.14	—	—	+.13	
Yes (95)	.176	24%	30%	13%	8%	32%	16%
No (46)	.146	8%	3%	6%	1%	6%	6%
PD Chronicity							
Beta 1	−.12	—	—	—	—	—	
Beta 2	−.13	—	—	—	—	—	
Yes (110)	.578	57%	74%	55%	37%	81%	67%
No (26)	.935	61%	59%	44%	47%	67%	67%

(*continued*)

TABLE 3—Continued

Time 1 (1981) Personality/Life Experience Predictors	Time 2 (1986) Problem Drinking Indicators						
	Average Consumption (ounces ethanol/day)	Heavy Episodic Drinking (% any)	Intoxication (% any)	Problem Consequences (% any)	Dependence Symptoms (% any)	Problem Drinking Index (% 1^{+})	Problem Drinker Sample Criteria (%1^{+})
Sexual Dysfunction, up to 1981							
NPD Onset							
Beta 1	—	—	—	—	—	—	
Beta 2	—	—	—	—	—	—	
Any (45)	.184	21%	22%	6%	5%	22%	10%
None (101)	.140	16%	19%	12%	5%	21%	12%
PD Chronicity							
Beta 1	—	—	+.19	+.19	+.18	+.29	
Beta 2	—	—	+.17	+.15	+.14	+.27	
Any (60)	.517	69%	81%	61%	52%	86%	75%
None (80)	.765	48%	59%	45%	29%	69%	59%
Heavy-Drinking Companions, 1981							
NPD Onset							
Beta 1	—	—	—	+.06	+.11	+.04	
Beta 2	—	—	—	+.07	+.11	+.01	
2^{+} (26)	.171	19%	25%	13%	0%	22%	13%
0–1 (130)	.157	17%	19%	9%	6%	21%	11%
PD Chronicity							
Beta 1	—	—	—	−.26	—	—	
Beta 2	—	—	—	−.20	—	—	
2^{+} (54)	.877	62%	73%	48%	37%	82%	65%
0–1 (88)	.550	56%	69%	56%	41%	76%	69%
Frequent-Drinking Partner, 1981							
NPD Onset							
Beta 1	—	—	—	—	−.17	—	
Beta 2	—	—	—	—	−.15	—	
Yes (22)	.171	19%	21%	9%	0%	21%	9%
No (135)	.157	17%	19%	10%	6%	21%	12%
PD Chronicity							
Beta 1	+.11	—	—	−.04	—	—	
Beta 2	+.14	—	—	−.07	—	—	
Yes (43)	.921	64%	69%	43%	39%	77%	62%
No (99)	.567	56%	71%	57%	40%	78%	70%

Note.—See Note, Table 2.

[a] In preliminary winnowing analyses, lifetime drug use was a stronger predictor of onset while current (1981) drug use was a stronger predictor of chronicity. Accordingly, only the lifetime use variable was included in the final regression equations for NPDs, and only the current use variable was used for PDs.

[b] Means and percentage rates are not provided for predictor categories with unweighted Ns less than 5.

predictors, only one—nontraditional feminine values—failed to predict any Time 2 drinking variables within either of the two subsamples.

Two variables often thought to be antecedents of problem drinking[16,17]—divorce or separation, and the "empty nest"—appeared to occur primarily after heavy or problem drinking in the present longitudinal bivariate analyses. The experience of a recent "empty nest" (last child leaving home in the past three years) did not predict Time 2 drinking behavior in either subsample, but among PDs, higher average consumption at Time 1 increased the likelihood that a respondent's last child would leave home during the three years prior to Time 2. Divorce and separation also occurred after rather than before heavy or problem drinking. Among NPDs, higher average consumption, heavy episodic drinking, and intoxication at Time 1 increased the likelihood of divorce or separation during the follow-up interval; similarly, among PDs, higher average consumption, intoxication, and PDI scores at Time 1 increased the chances of divorce or separation by Time 2. Contrary to expectations, being divorced or separated at Time 1 predicted lower levels of intoxication and lower PDI scores at Time 2 among PDs. We discuss additional bivariate relationships later in this paper, as an aid to interpreting multivariate findings.

Multivariate Prediction of Time 2 Drinking

Because some predictor variables were intercorrelated, it was important to examine the independent effects of each predictor, controlling for effects of all the others. Tables 2, 3, and 4 present the results of 12 unique hierarchical multiple regression equations predicting onset of problem drinking indicators among 1981 NPDs and chronicity of problem drinking indicators among 1981 PDs, using all predictors of a given Time 2 drinking variable surviving the winnowing process described above. Table 2 presents findings for Time 1 sociodemographic predictors; Table 3, findings for Time 1 personality and life-experience predictors; and Table 4, findings for all interim predictors. Although separated into three tables for convenience, the regression coefficients reflect the strength of each predictor when adjusted for *all* other predictors included in a given equation. Tables 2–4 also provide unweighted Ns, weighted means of

TABLE 4—Interim (1981–86) Sociodemographic, Personality, and Life Experience Predictors of Onset and Chronicity of Problem Drinking Indicators: Hierarchical Multiple Regression Coefficients with Unadjusted (Zero-Order) Weighted Means and Rates

	Time 2 (1986) Problem Drinking Indicators						
1981–86 Interim Predictors	Average Consumption (ounces ethanol/day)	Heavy Episodic Drinking (% any)	Intoxication (% any)	Problem Consequences (% any)	Dependence Symptoms (% any)	Problem Drinking Index (% 1⁺)	Problem Drinker Sample Criteria (%1⁺)
Became Unemployed, 1981–86							
NPD Onset							
Beta 2	—	+.08	+.12	—	—	—	
Yes (13)	.178	43%	68%	10%	0%	66%	10%
No (144)	.157	16%	16%	9%	5%	18%	11%
PD Chronicity							
Beta 2	+.26	—	—	—	—	—	
Yes (24)	.832	50%	82%	58%	33%	88%	67%
No (119)	.624	59%	68%	52%	41%	75%	67%
Became Employed Part time, 1981–86							
NPD Onset							
Beta 2	—	—	—	+.18	+.05	+.13	
Yes (34)	.225	22%	33%	27%	14%	33%	27%
No (123)	.143	17%	16%	9%	3%	18%	8%
PD Chronicity							
Beta 2	—	—	—	—	—	—	
Yes (30)	.572	63%	82%	48%	46%	84%	68%
No (113)	.683	56%	67%	55%	38%	75%	67%
Became Divorced/Separated, 1981–86							
NPD Onset							
Beta 2	—	—	+.02	—	+.01	+.01	
Yes (11)	.154	25%	44%	17%	17%	44%	25%
No (146)	.159	17%	17%	9%	4%	19%	10%
PD Chronicity							
Beta 2	—	—	−.15	—	−.24	−.15	
Yes (29)	1.028	48%	68%	53%	30%	72%	55%
No (114)	.582	60%	71%	53%	42%	78%	70%
Began Cohabiting, 1981–86							
NPD Onset							
Beta 2	+.18	—	+.16	—	—	+.18	
Yes (9)	.252	21%	46%	21%	6%	52%	21%
No (148)	.152	17%	18%	9%	5%	19%	11%
PD Chronicity							
Beta 2	+.13	—	+.10	—	—	—	
Yes (15)	1.104	58%	82%	65%	32%	82%	65%
No (128)	.613	58%	69%	52%	41%	77%	67%
Depressive Episodes, 1981–86							
NPD Onset							
Beta 2	—	—	—	+.23	+.16	+.12	
1^+ (40)	.155	24%	27%	21%	13%	28%	25%
None (115)	.156	15%	17%	5%	2%	19%	6%
PD Chronicity							
Beta 2	—	+.10	+.19	+.08	+.24	+.10	
1^+ (50)	.570	63%	81%	52%	54%	83%	64%
None (92)	.700	55%	64%	53%	32%	74%	68%
No. of Drugs Used, 1981–86							
NPD Onset							
Beta 2	—	+.08	—	—	+.07	—	
1^+ (65)	.164	25%	20%	8%	8%	21%	11%
None (90)	.160	14%	20%	10%	3%	22%	12%
PD Chronicity							
Beta 2	—	—	—	+.12	—	+.05	
1^+ (77)	.632	57%	75%	60%	45%	85%	72%
None (65)	.698	60%	64%	46%	34%	69%	63%
Premature Baby, 1981–86							
NPD Onset							
Beta 2	—	−.17	−.13	—	—	—	
Yes (2)[a]							
No (155)	.158	18%	20%	9%	5%	21%	11%

(continued)

TABLE 4—Continued

1981–86 Interim Predictors	Time 2 (1986) Problem Drinking Indicators: Average Consumption (ounces ethanol/day)	Heavy Episodic Drinking (% any)	Intoxication (% any)	Problem Consequences (% any)	Dependence Symptoms (% any)	Problem Drinking Index ($\% 1^+$)	Problem Drinker Sample Criteria ($\%1^+$)
PD Chronicity							—
Beta 2	—	—	—	−.04	−.18	—	
Yes (4)[a]							
No (139)	.670	58%	70%	55%	41%	77%	69%
Miscarriage/Stillbirth, 1981–86							
NPD Onset							
Beta 2	—	—	—	—	—	—	
Yes (1)[a]							
No (156)	.159	18%	20%	10%	5%	21%	11%
PD Chronicity							
Beta 2	—	—	—	−.12	—	—	
Yes (6)	.204	70%	70%	14%	28%	70%	28%
No (137)	.679	57%	70%	55%	40%	78%	69%
Infertility, 1981–86							
NPD Onset							
Beta 2	—	—	—	—	—	—	
Yes (6)	.205	22%	29%	35%	0%	47%	35%
No (151)	.156	17%	19%	9%	5%	20%	10%
PD Chronicity							
Beta 2	+.29	—	—	—	+.13	+.08	
Yes (15)	1.140	58%	84%	64%	69%	89%	69%
No (128)	.609	58%	69%	52%	37%	76%	67%
Sexual Dysfunction, 1981–86							
NPD Onset							
Beta 2	—	—	—	—	—	—	
Any (41)	.146	10%	21%	15%	7%	23%	16%
None (97)	.157	21%	20%	8%	4%	22%	10%
PD Chronicity							
Beta 2	—	−.16	—	+.10	—	—	
Any (43)	.650	45%	75%	57%	37%	81%	71%
None (93)	.623	67%	70%	52%	42%	76%	65%
1st-Step R^2, R^2 Change, and Total R^2 Values for Each Equation Predicting Onset among Nonproblem Drinkers							
1st-Step R^2	.02	.24***	.20***	.07	.09*	.13**	
R^2 Change	.03*	.04	.06	.09***	.04	.06	
Total R^2	.05	.28***	.25***	.16***	.13*	.19**	
1st-Step R^2, R^2 Change, and Total R^2 Values for Each Equation Predicting Chronicity among Problem Drinkers							
1st-Step R^2	.18***	.09*	.18**	.18*	.06	.20***	
R^2 Change	.14***	.03	.04	.05	.12***	.03	
Total R^2	.32***	.12*	.22**	.23*	.18***	.24***	

Note.—See Note, Table 2.

[a] Means and percentage rates are not provided for predictor categories with unweighted Ns less than 5.

(For R^2 values) *$p < .05$; **$p < .01$; ***$p < .001$.

average daily consumption, and weighted percentage rates on each of the Time 2 drinking measures for each predictor's primary category of effect. These unadjusted means and percentages from simple bivariate time-lagged comparisons are presented for descriptive purposes. However, for brevity, we limit our discussion to the more rigorous and conservative multiple regression analyses.

The numerous relationships summarized in Tables 2–4 mean that some relationships will be statistically significant simply by chance. For this reason, our discussion focuses primarily on those variables that predicted more than one Time 2 drinking variable, and on variables that showed consistent relationships to drinking within both NPD and PD samples or for both Time 1 and interim measures.

Time 1 vs. Interim Variables: Tables 2–4 show that both longer term (Time 1) and more recent (interim) influences are important for the longitudinal prediction of women's drinking. However, the time-order of interim variables and drinking changes cannot be specified precisely with the data analyzed here. Interim variables are most likely to be antecedents rather than consequences of drinking if they 1) are also significant as Time 1 predictors, and/or 2) are not consequences of increased drinking or drinking problems in the two-way bivariate analyses reported earlier.

Predictors of Onset: As shown in Table 1, of 153 drinking women with no indicators of problem drinking in 1981, by 1986 17 (11 percent) had developed at least one indicator, and seven (4 percent) reported two or three indicators. Tables 2–4 show that the most consistent predictors of onset of heavier drinking or adverse drinking consequences were younger age, cohabiting at Time 1 or during the follow-up interval, lifetime use of drugs other than alcohol, and depressive episodes during the follow-up interval. Cohabiting was significant as both a Time 1 and an interim predictor and predicted the largest number of Time 2 drinking variables (average consumption, heavy episodic drinking, intoxication, and PDI). Younger age predicted increased intoxication, and in equations without interim variables also predicted increased heavy episodic drinking and PDI scores, while the number of drugs used up to 1981 predicted increased intoxication and PDI scores and, in an equation without interim variables, increased alcohol dependence symptoms. Two Time 1 predictors—nontraditional sexual behavior and low self-esteem—each increased the likelihood of heavy episodic drinking at Time 2.

Predictors of Chronicity: Of the 140 1981 PDs for whom complete data are available, 51 (36.5 percent) continued to report two or three problem drinking indicators in 1986, while 43 (31 percent) reported only one indicator and 46 (33 percent) no longer reported any indicators. As shown in Table 3, the most consistent predictor of persistent problem drinking was Time 1 sexual dysfunction, which predicted Time 2 intoxication, drinking problems, alcohol dependence symptoms, and PDI scores. Depressive episodes during the follow-up interval predicted continued intoxication and dependence symptoms, and in an equation without interim variables, Time 1 depression predicted continued drinking problems.

Five other variables predicted that PDs would drink more heavily at Time 2: interim unemployment and infertility, and Time 1 part-time employment, nontraditional gender traits, and frequent-drinking partner. PDs who had never married by Time 1 were more likely to report Time 2 drinking problems and (in equations excluding interim variables) to report more Time 2 intoxication and higher PDI scores. Use of drugs other than alcohol at Time 1 increased the likelihood of continued alcohol dependence symptoms and (in the equation excluding interim variables) heavier drinking.

Predictors of Remission: Although the analyses reported here were not designed to study remission of problem drinking, some variables that were expected to predict increased drinking had opposite effects in the longitudinal analyses (indicated by negative beta coefficients in Tables 2–4). For the most part, these were isolated effects, involving only single dependent variables. Consistent with the bivariate effects of Time 1 divorce, which reduced Time 2 intoxication and PDI scores among problem drinkers, becoming divorced or separated during the follow-up interval decreased the level of Time 2 alcohol dependence symptoms reported by problem drinkers.

Discussion

Patterns of Change

As in earlier reports based on retrospective[11] and longitudinal[14,18] data, women's drinking behavior in this study showed considerable change over a five-year period. Eleven percent of women with no signs of problem drinking in 1981 reported at least one problem drinking indicator in 1986, and 33 percent of the 1981 problem drinkers were completely free of problems five years later. This fluctuation may in part reflect the relatively mild definition of problem drinking used in this study, in contrast to more stable and interrelated alcohol problems associated with severe problem drinking or clinical alcohol dependence. However, our findings may also indicate that frequently drinking problems among women in the general, nonclinical population are sporadic or intermittent, possibly related to shifts in roles, contexts, and other circumstances in women's lives.

Age as a Mediator of Drinking Change

Women's problem drinking appears to change more in youth than later in life. Younger age significantly predicted the onset of intoxication among NPDs and significantly predicted reductions in average consumption among PDs. Women age 21–34 were more likely than older women to show both onset and remission of problem drinking indicators, while chronicity of problem drinking was greatest among women age 35–49.

These patterns are consistent with previous findings of higher rates of heavy episodic drinking, intoxication, and drinking problems among younger than among older women;[4] with data showing greater remission of drinking problems among younger than among older men;[19] and with findings of greater chronicity among middle-aged than among younger or older problem drinking women.[20] Younger women's greater movement both into and out of risky drinking behavior may relate to fluctuations in drinking contexts and drinking partners within this age group, as well as to transitions in employment, marital, and childbearing roles.[10,21]

Nontraditional Life-style and Onset

Nonproblem drinking women who reported nontraditional sexual behavior (premarital sexual relations and/or masturbation before age 21) at Time 1 and those who cohabited at Time 1 were more likely than other NPDs to report heavy episodic drinking at Time 2, and those beginning cohabitation during the follow-up interval were more likely to report other problem drinking indicators. Nontraditional sexual behavior and cohabitation may be indicators of a life-style associated with greater freedom from the traditional moral constraints that have restricted both sexual and drinking behavior, especially in women.[9,22] This greater freedom from conservative moral constraints may carry with it increased drinking opportunities and more permissive drinking norms that can heighten the risks of excessive drinking behavior. In addition, premarital and cohabiting relationships, lacking institutional protection or support, may involve stressful uncertainties that can lead to stress-medicating alcohol consumption.

Sexual Dysfunction and Chronicity

The best single predictor of chronicity was sexual dysfunction, which predicted continued intoxication, drinking problems, alcohol dependence symptoms, and PDI scores. This finding may reflect a tendency of problem drinking women to self-medicate sexual difficulties with alcohol, given prevailing cultural beliefs that drinking reduces sexual inhibitions and facilitates sexual pleasure.[9,23,24] Since heavy alcohol consumption has detrimental effects on physiological sexual functioning,[25] a self-reinforcing cycle may occur in which heavy drinking becomes both cause and consequence of sexual dysfunction. Our findings suggest a need to address issues of sexuality and sexual dysfunction in the treatment and recovery of alcoholic women, and a need to assess drinking behavior among women presenting with sexual dysfunction—needs also noted by recent clinical writers.[26–28] Because recent research on sexuality and drinking has focused mainly on women, it will be important to determine in future research the extent to which sexual experience and sexual dysfunction also influence drinking and problem drinking in men.

Role Deprivation and Women's Drinking

Other predictors of chronicity included never having been married, being employed part time at Time 1, and becoming unemployed during the follow-up interval. Part-time employment during the follow-up interval also predicted onset of drinking problems among 1981 NPDs, as did cohabiting. We have suggested elsewhere that a lack or loss of stable, full-time employment or marital roles (e.g., being unmarried, unemployed, or employed part time) may constitute "role deprivation."[10] Role deprivation may increase the risk of alcohol abuse by reducing women's self-esteem and feelings of social worth, and by reducing contact with role partners who could provide social support and feedback about excessive drinking behavior. Role deprivation as a risk factor for the onset and chronicity of problem drinking in women contrasts with earlier hypotheses regarding role conflict or role overload as causes of women's problem drinking[29][†] but is consistent with other evidence linking multiple roles to better health outcomes.[30,31]

Other Drug Use and Women's Drinking

In preliminary bivariate analyses, different Time 1 drug use variables predicted Time 2 drinking for NPDs as compared with PDs: the number of different types of drugs used *up to* 1981 predicted Time 2 alcohol dependence symptoms and PDI scores among NPDs, while the number of different types of drugs *currently* used in 1981 predicted Time 2 consumption, dependence symptoms, and PDI scores among PDs. Similarly, in the multivariate analyses reported in Table 3, drug use up to 1981 increased the likelihood of subsequent intoxication and higher PDI scores (as well as dependence symptoms in the equation omitting interim variables) among NPDs, while current drug use in 1981 predicted continued dependence symptoms (and, in the equation without interim variables, higher average consumption) at Time 2 among PDs.

These patterns suggest that the relationships between women's drinking and their use of other drugs may differ at different stages of problem drinking. The onset of problem drinking may be facilitated by a woman's long-term tendency to use various psychoactive drugs—perhaps in an alternating pattern over time—to feel better, to have a good time, or to cope with problems.[8] Once drinking problems have developed, the *concurrent* use of alcohol and other drugs may have synergistic effects—both physiological and behavioral—that increase the risks of continued or worsened alcohol abuse.

Unexpected Findings

From previous cross-sectional findings[10,29] and clinical-retrospective reports,[16,17] we expected that divorce or separation and children's departure from home would predict the onset and/or chronicity of women's problem drinking. However, in the longitudinal bivariate analyses both of these experiences were more likely to follow than to precede problem drinking, and in multivariate analyses, becoming divorced or separated during the follow-up interval reduced the risk of Time 2 alcohol dependence symptoms among PDs. Causes and effects may be more complex than the simple time sequences here can reveal. For example, it is possible that a troubled marriage (perhaps involving sexual dysfunction or a problem-drinking spouse) may lead a woman into problem drinking, which then remits when the marital distress is relieved by divorce or separation. A more refined and focused analysis is needed to distinguish the effects of two different types of role relationships: positive, gratifying role relationships, whose loss may lead to increased drinking, and negative, frustrating role relationships, whose termination may result in reduced distress and decreased drinking.

Different Predictors of Different Drinking Behaviors

In the results presented here, different Time 1 and interim variables predicted different Time 2 drinking variables. Furthermore, these links between predictor and drinking variables were sometimes different for PDs and NPDs. We are clearly not dealing with a unitary entity—"problem drinking"— but rather with multiple aspects of risky drinking behavior and drinking consequences, which may be influenced differently by different characteristics of the drinker and her social environment.

Despite this complexity, a few generalizations are possible. Predictors of drinking patterns (average consumption, heavy episodic drinking, and intoxication) are more often situational variables (e.g., cohabiting, unemployment, frequent-drinking partner) or life-style variables (e.g., nontraditional sexual behavior, nontraditional gender traits), whereas predictors of adverse drinking consequences (drinking problems, alcohol dependence symptoms) include more individual psychological variables (in particular, depression and sexual dysfunction). Psychological variables are also somewhat more prominent predictions of chronicity than of onset. These patterns suggest that the factors that lead a woman into high-risk drinking are not the same as the factors that sustain or exacerbate problem drinking. Participation in a nontraditional lifestyle, involving greater freedom from conventional moral constraints on women's drinking and sexual behavior, may entail both drinking opportunities and stressful uncertainties that can increase a woman's risk of hazardous drinking behavior. Whether problematic drinking behavior continues may depend on other factors, including the extent to which a woman uses alcohol to deal with personal or interpersonal difficulties such as sexual dysfunction, depression, or role deprivation. Delineating more fully the predictors of change at these different stages of problem drinking, for specific subgroups of women (e.g., by age), would allow primary and secondary prevention efforts to be targeted more precisely to the characteristics of the drinker, her social environment, and the extent to which she has progressed into problem drinking behavior. ☐

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APPENDIX— Personality, Sexuality, and Drinking Variables*

Personality

Depressive Episodes, up to 1981: Number of depressive episodes (two or more weeks of depressed mood and three or more depressive symptoms, occurring within the same month; from Diagnostic Interview Schedule[32]) up to 1981 (0, 1, $2^\dagger$).

Depressive Episodes, 1981–86: Presence of any depressive episodes (DIS criteria) in the five years before 1986 survey (yes/no).

Anxiety, 1981: Whether respondent considers self a nervous or anxious person at time of 1981 survey (yes/no).

Anxiety, 1986: Whether respondent considers self a nervous or anxious person at time of 1986 survey (yes/no).

Low Self-Esteem, 1981: With reference to 1981 self-description on seven gender-typed traits (Personal Attributes Questionnaire[33]), in how many ways respondent would like to be different (two levels: some or many ways vs. a few or none).

Low Self-Esteem, 1986: As measured in 1981 (four levels).

Nontraditional Gender Traits, 1981: Nine-level index based on two traditionally feminine traits—devotion to others, understanding of others (Personal Attributes Questionnaire[33])—dichotomized at most nontraditional quartile.

Nontraditional Gender Values, 1981: Eleven-level index based on two traditionally feminine values—marriage and parenthood[34]—dichotomized at most nontraditional quartile.

Nontraditional Gender Values, 1986: As measured in 1981.

Sexuality

Liberal Sexual Morality, 1981: Moral approval of adult premarital sex with love, in 1981 (four levels).[22]

Liberal Sexual Morality, 1986: As measured in 1981.

Nontraditional Sexual Behavior up to 1981: Index based on having had premarital sexual relations and/or a self-induced sexual climax before age 21, reported in 1981 (three levels). The index label "Nontraditional Sexual Behavior" refers to traditional sociocultural norms disapproving of both masturbation and premarital sex (particularly for women), and is not meant to imply that either behavior is presently nonconventional (i.e., exhibited only by a "deviant" minority of women). There is considerable evidence that traditional norms of this sort can continue to affect behavior and associated moral judgments even when behavioral violations and attitudinal exceptions achieve a modal, perhaps even a majority, level in a statistical sense.[35]

Sexual Dysfunction up to 1981: Index based on having experienced, up to 1981, lifetime lack of sexual interest, vaginismus, lifetime lack of orgasm, or unreliable (<50 percent of occasions) orgasm (four levels).[9] Questions based on Kaplan's classification of female sexual dysfunctions;[36,37] some questions adapted from Diagnostic Interview Schedule.[32] Male respondents answered parallel questions about male sexual dysfunctions.

Sexual Dysfunction, 1981–86: Index based on having experienced, between 1981 and 1986, lifetime lack of sexual interest, vaginismus, lifetime lack of orgasm, or unreliable (<50 percent of occasions) orgasm (four levels).

Drinking and Drinking Consequences

Average Consumption: Mean ounces of ethanol per day in last 30 days, estimated by calculating separately and summing for wine, beer, and liquor:

- (usual number of drinking days per month) × (usual number of drinks per typical drinking day) ×
- (size of drink: wine = 4 oz.; wine coolers = 5 oz. [1986]; beer = 12 oz. or 16 oz.; liquor = self-reported ounces per drink) ×
- (ethanol content: regular wine = 12 percent [1981], 11 percent [1986]; fortified wine = 18 percent; wine coolers = 12 percent [1986]; regular beer = 4 percent; malt liquor = 5 percent [1986]; low-calorie beer = 3 percent [1986]; low-alcohol beer = 2 percent [1986]; liquor = 41 percent [1981], 40 percent [1986])/30 days.

Heavy Episodic Drinking: Frequency, in past 12 months, of having six or more drinks of wine, beer, or liquor in a single day (three levels: never in past 12 months; one–three days or four–seven days in past 12 months; eight to 11 days in past 12 months to five or more days per week).

Intoxication: Frequency, in past 12 months, of drinking enough to feel drunk—where drinking "noticeably affected your thinking, talking, and behavior" (three levels: never in past 12 months; one–three days in past 12 months; four days in past 12 months to five or more days per week).

Problem Consequences: Number of drinking-related problems experienced in past 12 months, from a list of nine problems (see text, "Drinking Variables"; three levels: 0, 1, 2 or more).

Alcohol Dependence Symptoms: Number of alcohol dependence symptoms experienced in past 12 months, from a list of five symptoms (see text; three levels: 0, 1, 2 or more).

Problem Drinking Index (PDI): Index summing the occurrence in the past 12 months of 1) any intoxication, 2) any problem consequence, and 3) any alcohol dependence symptom (range: 0–3). Not to be confused with subsampling criteria for 1981 problem drinkers (at least two of the following: one or more ounces ethanol per day; one or more problem consequences in past 12 months; one or more alcohol dependence symptoms in past 12 months).

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*Definitions of other variables are available upon request from the authors.

Race and Weight Change in US Women: The Roles of Socioeconomic and Marital Status

ABSTRACT

Background. The prevalence of overweight among Black women in the US is higher than among White women, but the causes are unknown.

Methods. We examined the weight change for 514 Black and 2,770 White women who entered the first Health and Nutrition Examination Survey (1971–75) at ages 25–44 years and were weighed again a decade later. We used multivariate analyses to estimate the weight-change effects associated with race, family income, education, and marital change.

Results. The mean weight change was greater for the Black women than the White women. Adjustments for height, duration of follow-up, baseline body mass index, and multiple demographic, social, and behavioral variables did not reduce this mean Black-White difference.

Conclusions. Among US women, Black race is independently associated with a reduced likelihood of major weight loss, but not with major weight gain. Women at greatest risk of weight gain are those with education below college level, those entering marriage, and those with very low family income. (*Am J Public Health* 1991;81:319–323)

Henry S. Kahn, MD, David F. Williamson, MS, PhD, and Judy A. Stevens, MS, MPH

Introduction

Cross-sectional surveys among United States adults have demonstrated that Black women have a higher prevalence of overweight than White women.[1,2] Overweight in US women is directly associated with low socioeconomic status.[3-6] Despite the association between Black race and low income or low education, controlling for socioeconomic status has not eliminated the overweight prevalence differences between Black women and White women.[1]

Prevalence alone does not explain the reasons for these observed associations. To understand the origin of overweight, we wondered whether race or socioeconomic factors during adulthood were independently associated with a woman's likelihood of gaining weight and whether marital status was involved (since this factor appears to be associated both with weight status[5] and with race).

Rac[illegible]cioeconomic status, or marital stat[illegible]ht contribute to overweight prevalence either by promoting weight gain or by inhibiting weight loss. To clarify the circumstances that led to the current weight distribution, we analyzed the mean weight changes, the major weight gains, and the major weight losses of a nationally representative prospective cohort of adult US women that was followed up over approximately a 10-year interval.

Methods

Study Population

This report analyzes public-use computer tapes containing the data from the First National Health and Nutrition Examination Survey[7,8] (NHANES-I) and the NHANES-I Epidemiologic Follow-up Study.[9] The subjects of NHANES-I, a probability sample of the US civilian, noninstitutionalized population resident in the contiguous 48 states, were initially interviewed and weighed in 1971–75. We limited our analysis to those who entered at ages 25–44 because these younger adult subjects had a greater risk of weight gain than older women.[10]

The initial NHANES-I population sample used for this analysis included 3,589 White women and 735 Black women between the ages of 25 and 44. (Seventy-two women in this age range with other racial designations are not included in this report.) Between 1982 and 1984 the NHANES-I Epidemiologic Follow-up Study successfully traced 3,210 (89 percent) of these White women and 621 (84 percent) of these Black women.[11] We excluded from this cohort women who had died since their initial examinations (50 Whites, 33 Blacks) and women for whom a valid, nonpregnant weight was not available at either baseline or follow-up (390 Whites, 74 Blacks). Thus, the 3,284 subjects included in this study of weight change represent 77 percent of the White women and 70 percent of the Black women initially enrolled in the NHANES-I cohort.

The mean duration of the follow-up interval was 10.1 ± 1.2 (SD) years.

From the Center for Chronic Disease Prevention and Health Promotion, Centers for Disease Control, and the Department of Community and Preventive Medicine, Emory University School of Medicine. Address reprint requests to David F. Williamson, MS, PhD, Division of Nutrition, Mail Stop A41, Center for Chronic Disease Prevention and Health Promotion, CDC, 1600 Clifton Road, Atlanta, GA 30333. This paper, submitted to the Journal February 5, 1990, was revised and accepted for publication October 23, 1990.

Definitions of Weight Change

Baseline measurements of height and weight were obtained under controlled conditions, with the subject wearing only disposable paper uniforms and foam rubber slippers. The follow-up weight was obtained in the subject's home after she removed her shoes and articles of heavy clothing. These follow-up values were corrected for the weight of the subject's indoor clothing by subtracting 1.6 kg from the measured weight.[10]

For each subject, weight change was categorized as either *major weight gain* (an increase of 13.0 kgs or more), *major weight loss* (a decrease of 7.0 kgs or more), or *mid-range weight change* (weight changes > −7 kgs but < +13 kgs).

Statistical Analysis

Considering weight change as a continuous variable, we used multiple linear regression models to estimate the effects associated with race, family income, education, and marital change. Our simple model adjusted each of these variables only for age, height, and duration of follow-up. Our fully adjusted model included adjustments for race, family income, education, marital change, age, height, duration of follow-up, body mass index (BMI, kg/m^2) at baseline, smoking, physical activity, parity, and rural/urban background (see Appendix for variable definitions).

Considering weight change as a categorical variable, we used multiple logistic regression analysis (employing the SAS procedure PROC LOGIST[12]) to estimate the odds ratio of either major weight gain or major weight loss for the various categories of race, family income, education, and marital change. Simple and fully adjusted models were developed as described above. For these analyses, the mid-range outcome of weight change was considered the "non-disease" group. Logistic regression analyses excluded 19 subjects with various unknown values.

We also performed multiple linear and logistic regression analyses stratified by overweight status at baseline. The overweight group at baseline (BMI ≥ 27.3) included 221 Blacks and 547 Whites; the group not overweight at baseline (BMI < 27.3) included 293 Blacks and 2,223 Whites.

Results

Mean Changes in Weight

The mean (±SD) change in weight was greater for Black women (3.8 ± 9.6 kgs) than for White women (2.8 ± 8.0) (difference = 1.0 kgs [95% confidence interval = 0.1, 1.9]). Blacks had a significantly wider variance of this weight-change variable (p = .0001, F-test for inequality of the variances). Adjustments for age, height, duration of follow-up, and other variables (Table 1) did not reduce this Black-White difference in mean weight change. Fully adjusted analyses stratified by overweight status at baseline showed that the Blacks' excess mean weight gain was 3.7 kgs (95% CI = 1.7, 5.8) among women who were initially overweight; among women not initially overweight, the Blacks' excess was only 0.8 kg (95% CI = −.1, 1.6).

The fully adjusted multiple linear regression model (Table 1, right column) demonstrated that women with only a 12th grade education had significantly greater weight gain compared to women who had gone beyond the 12th grade. When compared with women who were consistently married, women who became married during the 10-year interval experienced a greater weight gain, and women whose marriage ended experienced a relative weight loss.

Risk Factors for Major Weight Gain

Black women experienced a higher crude incidence of major weight gain (13.4 percent) than did the White women (7.7 percent).

TABLE 1—Effect of Race, Family Income, Educational Attainment, and Marital Change on Mean Weight Change, by Multiple Linear Regression Modeling; US Women aged 25–44 at Baseline, NHANES-I Baseline (1971–75) and NHANES-I Follow-up (1982–84) [n = 3284]

	Increment in the Mean Weight Change in Kilograms (95% confidence interval)			
Variables	Adjusted for Age, Height, and Duration of Follow-up		Fully Adjusted Model*	
Race				
White	0.0	—	0.0	—
Black	+1.0	(+.2, +1.8)	+1.7	(+0.8, +2.5)
Family income, both interviews				
Adequate	0.0	—	0.0	—
Moderately low	−0.2	(−1.0, +0.7)	−0.1	(−0.9, +0.8)
Very low	−0.2	(−1.0, +0.6)	−0.4	(−1.3, +0.6)
Unknown	−0.1	(−1.1, +0.9)	−0.2	(−1.2, +0.8)
Education (highest grade)				
> 12th	0.0	—	0.0	—
12th	+0.6	(−0.1, +1.3)	+0.9	(+0.2, +1.6)
< 12th	+0.2	(−0.6, +1.0)	+0.7	(−0.1, +1.6)
Unknown	−1.0	(−4.7, +2.8)	+0.2	(−3.6, +4.0)
Marital change				
Consistently married	0.0	—	0.0	—
Marriage ended	−1.2	(−2.1, −0.3)	−1.0	(−1.9, −0.2)
Became married	+2.3	(+0.9, +3.8)	+2.1	(+0.6, +3.4)
Stayed unmarried	−0.1	(−0.8, +0.7)	+0.3	(−0.7, +1.2)
Unknown	−6.4	(−16, +2.9)	−5.4	(−15, +4.0)

*Includes race, family income, education, marital change, age, height, duration of follow-up, BMI at baseline, smoking, physical activity, parity, and rural/urban background.

TABLE 2—Effect of Race, Family Income, Educational Attainment, and Marital Change on the Risk of Major Weight Gain (MWG), by Multiple Logistic Regression; US Women Aged 25–44 at Baseline Who Experienced MWG (n = 281) or Mid-range Weight Change (n = 2,759) between NHANES-I Baseline (1971–75) and NHANES-I Follow-up (1982–84)

	Estimated Odds Ratios for MWG (95% confidence interval)			
Variables	Adjusted for Age, Height, and Duration of Follow-up		Fully Adjusted Model*	
Race				
White	1.00	—	1.00	—
Black	1.79	(1.33, 2.41)	1.21	(0.85, 1.71)
Family income				
Adequate	1.00	—	1.00	—
Moderately low	1.60	(1.12, 2.29)	1.38	(0.94, 2.04)
Very low	2.08	(1.52, 2.85)	1.71	(1.15, 2.55)
Unknown	1.16	(0.73, 1.87)	1.21	(0.74, 1.98)
Education (highest grade)				
> 12th	1.00	—	1.00	—
12th	1.46	(1.05, 2.04)	1.19	(0.84, 1.68)
< 12th	1.58	(1.10, 2.27)	0.91	(0.60, 1.39)
Unknown	1.98	(0.43, 9.07)	0.83	(0.17, 4.10)
Marital change				
Consistently married	1.00	—	1.00	—
Marriage ended	0.90	(0.60, 1.35)	0.80	(0.52, 1.23)
Became married	1.55	(0.92, 2.62)	1.53	(0.88, 2.65)
Stayed unmarried	1.28	(0.92, 1.78)	1.00	(0.65, 1.52)

*Includes race, family income, education, marital change, age, height, duration of follow-up, BMI at baseline, smoking, physical activity, parity and rural/urban background.

When the logistic regression model was adjusted only for age, height, and duration of follow-up (Table 2, left column), Black race was significantly associated with major weight gain. Both low family income and low education were also associated with major weight gain (Table 2, left column).

However, in the fully adjusted model (Table 2, right column), Black race was only weakly associated with major weight gain. Stratification by overweight status at baseline did not alter this negative finding. In the fully adjusted model the strongest risk factors associated with major weight gain were very low family income (OR = 1.71) and becoming married during the follow-up interval (OR = 1.53).

The risk of major weight gain was also independently associated with being overweight at baseline in our fully adjusted model. Compared with women who entered the cohort in the lowest BMI stratum (BMI < 19.1), those in successively higher strata of baseline BMI had odds ratios for major weight gain of 0.64 (95% CI = 0.38, 1.07), 1.32 (0.79, 2.19), 2.23 (1.30, 3.82), and 2.26 (1.27, 4.04). If the larger multivariate model excluded adjustment for baseline BMI, the association between Black race and major weight gain yielded OR = 1.41 (95% CI = 1.00, 1.99).

Risk Factors for Major Weight Loss

Black women also experienced a higher crude incidence of major weight loss (8.8 percent) than did the White women (6.5 percent). However, the major weight loss logistic regression model, adjusting for age, height, and duration of follow-up (Table 3, left column), demonstrated that the odds ratio for Black race was only 1.35 (95% CI = 0.95, 1.92). After fully adjusting for other variables (Table 3, right column) Black race had an inverse (protective) association with major weight loss (OR = 0.64). Fully adjusted analyses stratified by baseline BMI indicated that this inverse association was limited to women who were overweight at baseline (OR = 0.53 [95% CI = 0.32, 0.89]) and was not observed among women who were not initially overweight (OR = 1.08 [95% CI = 0.51, 2.30]).

Very low family income was directly associated with the risk of major weight loss in the fully adjusted model (OR = 1.86).

Discussion

Several methodologic or interpretive issues should be considered when reviewing the weight-change experience of this large, representative, prospective cohort.

Approximately one-fourth of the population originally enrolled in NHANES-I was omitted from the analyses because of loss to follow-up or lack of a valid weight measurement. The effect of their missing data is unknown. However, any substantial bias would most likely have been due to diseases which might have caused these women both to be missing from the cohort and to have gained either more or less weight than those who participated fully. For these missing, hypothetically diseased, women to have changed substantially our results, they would have had to experience an extraordinary degree of consistently increasing weight (e.g. edema or ascites) or consistently decreasing weight (e.g. cachexia). This is unlikely since the enrolled cohort was relatively young and during the 10-year interval only a small proportion would have been expected to acquire a major, chronic disease. Systematic biases due to other causes seem even less likely.

Our principal outcome events, major weight gains and losses, were determined by the difference between weight measurements made at only two points widely separated in time. Fluctuations in weight that may have occurred during the decade of follow-up are not reflected in our data. Hence our case definitions for both major weight gain and major weight loss probably failed to include everyone who had a major weight change between baseline and follow-up exams. Semi-annual weights obtained in a four-year Dutch study suggest that overweight women and women with the highest rates of weight gain or weight loss are also the women who experience the widest weight fluctu-

TABLE 3—Effect of Race, Family Income, Educational Attainment, and Marital Change on the Risk of Major Weight Loss (MWL), by Multiple Logistic Regression; US Women Aged 25–44 at Baseline Who Experienced MWL (n = 225) or Mid-range Weight Change (n = 2,759) between NHANES-I Baseline (1971–75) and NHANES-I Follow-up (1982–84)

	Estimated Odds Ratios for MWL (95% confidence interval)			
Variables	Adjusted for Age, Height, and Duration of Follow-up		Fully Adjusted Model*	
Race				
White	1.00	—	1.00	—
Black	1.35	(0.95, 1.92)	0.64	(0.41, 0.97)
Family income, both interviews				
Adequate	1.00	—	1.00	—
Moderately low	1.69	(1.13, 2.52)	1.25	(0.79, 1.97)
Very low	2.43	(1.73, 3.42)	1.86	(1.18, 2.95)
Unknown	1.41	(0.86, 2.29)	1.73	(1.01, 2.96)
Education (highest grade)				
> 12th	1.00	—	1.00	—
12th	1.45	(0.99, 2.12)	0.92	(0.61, 1.41)
< 12th	1.86	(1.24, 2.78)	0.69	(0.42, 1.12)
Unknown	2.76	(0.60, 12.8)	0.97	(0.17, 5.51)
Marital change				
Consistently married	1.00	—	1.00	—
Marriage ended	1.57	(1.08, 2.38)	1.18	(0.76, 1.83)
Became married	0.23	(0.06, 0.93)	0.25	(0.06, 1.05)
Stayed unmarried	1.16	(0.80, 1.69)	0.74	(0.45, 1.21)

*Includes race, family income, education, marital change, age, height, duration of follow-up, BMI at baseline, smoking, physical activity, parity, and rural/urban background.

ations.[13] Thus, our observed incidence rates for major weight change likely represent low estimates for the occurrence of major weight gain or major weight loss over a 10-year period.

After adjusting for multiple variables, Black women had a greater increase in mean weight compared to White women in this cohort (Table 1). This finding is consistent with US prevalence data. However, after full multivariate adjustments we found that Black women were not significantly more likely than White women to experience a major weight gain (Table 2). The racial effect on mean weight change could readily be viewed as a result of the Black women's decreased tendency to lose weight (Table 3). Our analyses stratified by baseline overweight status demonstrated that it was only among women who were overweight at baseline that the Blacks showed a clear excess of mean weight gain and a protective effect against major weight loss.

This reduced likelihood of weight loss among overweight Black women deserves further study. It could reflect acquired cultural factors (e.g. specific dietary exposures, perceptions of ideal body size) that may have made Black women less able or less motivated than White women to lose weight. Genetic factors may also play a role. Our data provide no basis for choosing among these explanations. As with any studies that collect data on race,[14] the NHANES-I criterion for being Black (or White) may have historical, social, or cultural significance that exceeds any clear genetic meaning.

We noted that the risks of both major weight gain and major weight loss in the NHANES-I follow-up cohort were directly associated with high BMI at baseline. Similar observations have been made previously in other female populations from the US[15] and Europe.[13,16] Since the Black women in our study were more overweight at baseline than were the White women, preexisting weight status was a potential confounder of the relationship between women's weight change and race. If the baseline BMI was not taken into account, multiple adjustments for current social, demographic, and behavioral factors reduced the Blacks' odds ratio for major weight gain from 1.79 to 1.41. Including adjustment for baseline BMI (along with the other variables) further reduced the major weight gain odds ratio to 1.21. It is possible that our adjustment for baseline BMI served as a surrogate adjustment for the prior weight-related experiences of cohort members before they entered the follow-up study. If low income and low education are associated with increased mean or major weight gains, then these circumstances common among US Black women before they entered NHANES-I may have explained their having baseline BMIs that were higher than those of the White women.

The serious health risks associated with adult overweight and the low success rates associated with its treatment suggest that we must develop a public health strategy for the prevention of weight gain among adults. Our findings indicate that women in greatest need of help in preventing weight gain are those with education below college level, those with low family incomes, and those who become married. Black race does not appear to be an independent adult risk factor for weight gain. Future research should address the mechanisms by which the diet and social stresses of women with low socioeconomic status may influence their weight changes. Studies of eating practices and physical activity changes associated with entering marriage might likewise contribute to the prevention of adult weight gain. ☐

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APPENDIX

Definitions of Other Variables

Race was classified by interviewer observation as "White, Negro or other"; if the appropriate category could not be marked by observation, then the respondent was asked her race. (Hispanics not regarded as "Negro" were included with "White" unless definitely known to be American Indian or other non-White race, in which case they were excluded from this analysis.)

The baseline interview also categorized the subject's *educational level* as the "highest grade of regular school ever attended."

The *family income* variable used categorical interview information collected both at NHANES-I baseline and at follow-up to classify women as having very low, moderately low, or adequate levels of family income across the ten-year interval:

	Family Income at Follow-up (1982–84)		
	< $15,000	15,000–$19,999	$20,000+
Family Income at Baseline *(1971–75)*			
$10,000 +	Adequate	Adequate	Adequate
$7,000–$9,999	Mod. Low	Mod. Low	Adequate
< $7,000	Very Low	Mod. Low	Adequate

Based on contemporaneous median family incomes in the U.S.,[17] women in the "very low" income level had family incomes below about 59% of median family incomes for both time periods. Women in the "moderately low" income level had family incomes below about 82 percent of median family incomes (but not repeatedly below about 59 percent of median) for both time periods.

Marital change was defined as consistently married, marriage ended, became married, or stayed unmarried, depending on whether the subject reported being married or not at baseline and at follow-up. Women who might have become divorced and remarried during the follow-up interval were included within the "consistently married" category.

Age at baseline was stratified by five-year intervals.

Height and *duration of follow-up* were treated as continuous variables.

The subjects' *weight status at baseline,* as estimated by body mass index (BMI, kg/m^2), was divided into five categories with break points at BMIs of 19.1, 23.0, 27.3, and 32.3.

Among the additional variables considered as possible confounders, *smoking status* was defined at follow-up as never smoked, continuing smoker, and ex-smoker. Self-described recreational and nonrecreational *physical activity,* determined at both baseline and follow-up, was defined in three levels as consistently low, moderate or inconsistent, and consistently high. *Parity* at the time of follow-up was stratified as 0, 1-2, 3-4, and 5+. *Rural/urban background* was determined by the subject's response at follow-up to the question "Would you say that you have lived most of your life in a rural area, or the city or in the city suburbs?"

Gender Differences in Cigarette Smoking and Quitting in a Cohort of Young Adults

Phyllis L. Pirie, PhD, David M. Murray, PhD, and Russell V. Luepker, MD

ABSTRACT

Background. Smoking among young women is associated with a variety of negative health outcomes. Gender specific influences on smoking, quitting and attempting to quit are hypothesized to occur and may have implications for cessation programs.

Methods. Telephone surveys were conducted in a large (n = 6,711) cohort of young men and women (average age 19.2 years) which was first established in 1979 and has been resurveyed several times since then. Questions concerned smoking, successful and unsuccessful attempts to quit, withdrawal symptoms during quit attempts, and concerns about quitting.

Results. More women than men reported current smoking (26.5 vs 22.6 percent), but quitting attempts, successful and unsuccessful, were equally common. Withdrawal symptoms were reported equally, except for wanting to eat more than usual and weight gain, both of which were reported more often by women than men. Women smokers reported substantially more concern about weight gain if they quit smoking (57.9 vs 26.3 percent expressing concern).

Conclusions. Targeted programs are needed to address issues of concern to young women smokers, particularly fear of gaining weight. (*Am J Public Health* 1991;81:324–327)

Introduction

The prevalence of smoking and the burden of smoking-related illness in this country are beginning to shift toward women. Among men, cross-sectional surveys on smoking status have presented a picture of steadily declining smoking rates from the 1960s to the present, coupled with a steady increase in the number of ex-smokers.[1-3] Among women, the percentage decline in smoking prevalence has been smaller.[3,4] Smoking prevalence rates in the youngest groups of adult women are beginning to exceed those in men, at least among whites,[1,5,6] and overall smoking rates among women are projected to exceed those among men by the latter part of the 1990s.[7]

Smoking in young women is particularly troubling because it not only poses hazards to their own long-term health,[8] but has immediate impacts on reproductive function[9,10] and the health of their children.[11,12] With a large population of young women continuing to take up smoking, there is a clear need to learn more about ways to help them quit smoking, as well as barriers to their quitting. This paper is an exploration of smoking, quitting, and attempted quitting behavior in a population of young adults, with particular attention to the experiences of young women.

Methods

The young adults in our study originally comprised the entire seventh-grade enrollment of four suburban districts in the Minneapolis-St. Paul, Minnesota, metropolitan area in two subsequent years, 1979 and 1980. They took part in two adolescent smoking studies, which have been described elsewhere.[13-15] Briefly, as seventh-graders, these students were exposed to variations of a "social influences" smoking prevention curriculum; long-term follow-up demonstrated modest differences in smoking prevalence among students in the various conditions through about eleventh grade,[14] but no differences after that time.[15]

The cohort of participating students was reconstructed in 1984-85 and resurveyed annually through 1988.[16,17] The data reported here are from the 1986-87 follow-up survey, at which time the participants were one or two years out of high school. The response rate for this survey was 94.6 percent, representing 6,711 of a total possible 7,096 interviews. An additional 28 members of the initial cohort were deceased or severely disabled and thus unavailable for interview.

Participants were randomly assigned to one of 11 months (excluding December) for annual follow-up and interviewed by telephone between October 1986 and November 1987. Data collected during the interview included items on current and former smoking status, quit attempts, withdrawal symptoms, perceived barriers to quitting, perceived pressures to quit, current weight and concern about weight. The smoking status variables reported here were based on self-report. In a separate analysis, a sample of individuals provided saliva samples for validation of smoking status;[18] results indicated that telephone self-report data are largely accurate in this population, although recent

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From the Division of Epidemiology, University of Minnesota School of Public Health. Address reprint requests to Phyllis L. Pirie, PhD, Division of Epidemiology, School of Public Health, University of Minnesota, Minneapolis, MN 55455. This paper, submitted to the Journal March 27, 1990, was revised and accepted for publication September 19, 1990.

TABLE 1—Smoking and Quitting Behaviors in a Cohort of Young Men and Women with an Average Age of 19.2 Years

Behaviors	Men (95% CI) n = 3455	Women (95% CI) n = 3256
Of total		
% Ever smoked ≥ 100 cigarettes in lifetime	32.1 (30.5, 33.7)	37.6 (35.9, 39.3)
% Current smokers	22.6 (21.2, 24.0)	26.5 (25.0, 28.0)
% Ex-smokers	9.6 (8.6, 10.6)	11.2 (10.0, 12.2)
Of Current Smokers	(n = 779)	(n = 863)
Number of cigarettes smoked in the last 24 hours	16.2 (13.6, 18.8)	14.3 (12.0, 16.6)
% Tried to quit in past 12 months	62.6 (59.2, 66.0)	66.9 (63.8, 70.0)
% Would like to quit	76.1 (73.1, 79.1)	72.4 (69.4, 75.4)
% Changed brands in past 12 months to reduce tar and nicotine	32.4 (29.1, 35.7)	34.5 (31.3, 37.7)
Of those who have tried to quit in the last 12 months	(n = 487)	(n = 575)
Number of quit attempts lasting more than 1 week	1.6 (0.5, 2.7)	1.6 (0.6, 2.6)
Length of longest quit attempt (days)	28.5 (24.5, 32.5)	36.7 (32.8, 40.6)

TABLE 2—Self-Reported Smoking Habits Reflective of Addictive or Dependent Smoking Among Young Adult Current Smokers, by Gender

Smoking Habits	Men (95% CI) n = 779	Women (95% CI) n = 863
% Always inhale	93.6 (91.9, 95.3)	93.4 (91.7, 95.1)
% Smoke more in morning than in rest of day	13.4 (11.0, 15.8)	13.3 (11.0, 15.6)
% Smoke within 30 minutes of getting up	44.2 (40.7, 47.7)	35.6 (32.4, 38.8)
% First cigarette in morning is one would most hate to give up	32.3 (29.0, 35.6)	24.9 (22.0, 27.8)
% Find it hard not to smoke in places where it's forbidden	13.6 (11.2, 16.0)	11.6 (9.5, 13.7)
% Smoke even when ill	16.5 (13.9, 19.1)	12.6 (10.4, 14.8)

quitting may be overreported, a discrepancy unlikely to affect the gender-specific comparisons reported here.

Simple proportions and their 95 percent confidence intervals are reported for categorical variables, and means and their confidence intervals for continuous variables. All analyses were conducted using SAS.[19]

Results

The average age of the respondents in this survey was 19.2 years; 51 percent were male. Nearly all of the students were White; 37 percent were enrolled in four-year colleges and 23 percent in vocational schools or two-year colleges. All but 7 percent had graduated from high school.

As shown in Table 1, more women than men were current smokers, and more had ever smoked. The proportion who reported themselves as former smokers, however, was approximately the same in men and women. No gender differences were seen in amount of current smoking or on most of the questions designed to elicit quitting behaviors, although women self-reported slightly longer quit attempts.

The genders differed slightly on items reflecting addictive or dependent smoking[20] (Table 2) with more men reporting that they smoked within 30 minutes of getting up and that the first cigarette in the morning was the one they would most hate to give up.

Both current and former smokers were asked about withdrawal symptoms (Table 3), with unsuccessful quitters reporting higher withdrawal symptoms on all symptoms except weight gain. Women were more likely than men to report wanting to eat more than usual, and they also reported more weight gain. The suggestion that weight gain is differentially a problem for women (as compared to men) is not supported by the self-reported weights, however. Among women, the quitters weighed about the same as current smokers (127.2 lbs for each group, on average), while among men, the successful quitters were heavier than current smokers (167.3 vs 162.0 lbs).

Finally, we examined various beliefs about the social and physical aspects of quitting smoking which may act as barriers to quitting, as well as potential social pressures to quit (Table 4). Women were more likely to agree with two of the "barriers" items: "If I quit smoking, I would probably gain a lot of weight" and, "If I quit smoking, it would be hard to go out with friends who smoke." Women were also more likely to agree that they felt social pressure to quit, reporting that, "A doctor has strongly urged me to quit smoking" and, "The people I'm closest to would like me to quit smoking."

Discussion

Contrary to reports in the past that women smokers have had a much harder time quitting smoking than men,[20,21] these data indicate that both quitting attempts and successful quitting are equally common in both genders for young adults. This may be a positive indication that future smoking patterns will show similar trends for men and women, with higher quit rates among women than are seen in existing cross-sectional surveys of adult women. On the other hand, these data do indicate that more young women than young men are beginning to smoke and are continuing the habit into early adulthood, a time when the risks of smoking to their personal and reproductive health begin to be manifest.

It is interesting that the unsuccessful quitters reported higher levels of withdrawal symptoms on nearly all items, suggesting that elevated levels of withdrawal symptoms may have prevented their successful cessation, although the unsuccessful nature of the attempt may also have led them to overreport symptom severity in order to rationalize failure.

At least one study has found that women who attempt to quit smoking report more severe withdrawal symptoms than do men,[22] although others have failed to confirm this finding.[23,24] In the young adult cohort reported here, among both successful and unsuccessful quitters, women did not generally report more withdrawal symptoms than did men, in

TABLE 3—Percent of Young Adults Reporting Withdrawal Symptoms, by Gender:

Withdrawal Symptoms	Men (95% CI) n = 454	Women (95% CI) n = 538
*Current smokers—%**		
Irritability or impatience	54.8 (50.2, 59.4)	56.7 (52.5, 60.9)
Difficulty concentrating	28.2 (24.1, 32.3)	26.4 (22.7, 30.1)
Restlessness	57.7 (53.2, 62.2)	53.9 (49.7, 58.1)
Craving for a cigarette	82.8 (79.3, 86.3)	81.9 (78.6, 85.2)
Wanted to eat more than usual	31.9 (27.6, 36.2)	49.5 (45.3, 53.7)
Fatigue or lack of energy	19.0 (15.4, 22.6)	19.6 (16.2, 23.0)
Wanted to drink more fluids than usual	48.0 (43.4, 52.6)	41.5 (37.3, 45.7)
Weight gain	14.5 (11.3, 17.7)	26.1 (22.4, 29.8)
Former smokers—%†	(n = 229)	(n = 224)
Irritability or impatience	40.6 (34.2, 47.0)	33.0 (26.9, 39.2)
Difficulty concentrating	15.7 (11.0, 20.4)	8.9 (5.2, 12.7)
Restlessness	36.7 (30.4, 42.9)	29.9 (23.9, 35.9)
Craving for a cigarette	64.6 (58.4, 70.8)	62.5 (56.2, 68.8)
Wanted to eat more than usual	26.2 (20.5, 31.9)	47.8 (41.2, 54.3)
Fatigue or lack of energy	13.5 (9.1, 18.0)	15.2 (10.5, 19.9)
Wanted to drink more fluids than usual	35.4 (29.2, 41.6)	28.6 (22.6, 34.5)
Weight gain	19.2 (14.1, 24.3)	29.5 (23.5, 35.4)

**Current smokers who have attempted to quit within last 12 months; for most recent attempt lasting more than 24 hours.

†Former smokers, who quit within one year of interview.

TABLE 4—Percent Agreement with Statements Reflecting Barriers and Pressures to Quit Smoking Reported by Young Adult Current Smokers, by Gender

Barriers/Pressures	Men (95% CI) n = 779	Women (95% CI) n = 863
Barriers to Quitting—%		
If I quit smoking, I would probably gain a lot of weight	26.3 (23.2, 29.4)*	57.9 (54.6, 61.2)
If I quit smoking, I would feel tense and irritable	62.1 (58.7, 65.5)	67.7 (64.6, 70.8)
I enjoy smoking too much to quit	50.4 (46.9, 53.9)	50.9 (45.6, 54.2)
If I quit smoking, it would be hard to go out with friends who smoke	32.7 (29.4, 36.0)	46.5 (43.2, 49.8)
If I got tense and irritable from not smoking, it would make the people I'm closest to unhappy	54.7 (51.2, 58.2)	59.2 (55.9, 62.5)
Pressures to Quit—%		
A doctor has strongly urged me to quit smoking	20.2 (17.4, 23.0)	31.5 (28.4, 34.6)
The people I'm closest to would like me to quit smoking	63.5 (60.1, 66.9)	70.6 (67.6, 73.6)

*95% confidence interval.

fact, among the successful quitters, men tended to report more symptoms. Weight concerns, however, emerged in opposition to this general pattern, with both successful and unsuccessful quitters among women reporting eating more and gaining weight during quit attempts. Current women smokers strongly endorsed the notion that they would gain weight if they quit smoking. Interestingly, although women reported more concern about weight, the evidence did not suggest any large weight gains among the successful quitters in this group, who weighed, on average, exactly as much as their smoking counterparts. Although several published studies indicated that weight gain after quitting smoking was greater in women than men,[1] that finding was not supported in this population. It was clear, however, that concern about gaining weight is much more common in women than in men, and that such concern was excessive in light of the magnitude of the problem. While 57.9 percent of women smokers in our population agreed that they would probably gain a lot of weight if they quit smoking, only 26.2 percent of quit attempters and 29.5 percent of actual quitters reported gaining weight. Even those gains must have been modest or readily controlled, as average weights for quitters were not higher than for current smokers.

One limitation of this study is the nonrepresentative nature of the sample, which consists primarily of White young adults, the majority of whom are pursuing education beyond high school. This is reflected in the self-reported smoking rates, which are lower than those reported from national samples.[4]

The data reported here provide support for the observation that young women are smoking cigarettes in greater numbers than young men, and concern about weight gain emerges as a major issue for these young women smokers. Given the risks to their familes as well as to themselves, public health cessation efforts targeted to women continue to be warranted. Targeted programs need to address issues of concern to women, particularly fear of gaining weight. ☐

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The Food and Nutrition Section, American Public Health Association, is announcing the availability of the Helen R. Stacey and Joseph A. Walsh Awards for Graduate Education at the Master's Degree level in Public Health Nutrition. Three awards are given consisting of $2,000 each. These awards must be used for the academic year 1991–92. *Applications must be submitted no later than May 1, 1991.*

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Intake of Tapwater and Total Water by Pregnant and Lactating Women

Abby G. Ershow, ScD, Linda M. Brown, MPH, and Kenneth P. Cantor, PhD

ABSTRACT

Background. Despite theoretically higher requirements for water due to physiologic demands of pregnancy and lactation, little is known of actual ranges of intake in pregnant and lactating women.

Methods. Population-based estimates of total water and tapwater intake in women of reproductive age were derived using data from the 1977–78 USDA Nationwide Food Consumption Survey. Three-day average intakes were calculated for 188 pregnant women, 77 lactating women, and 6,201 non-pregnant, non-lactating control women.

Results. Total water intake (mean ± SD) was 1,940 ± 686 g/day (median 1,835) for control women, 2,076 ± 743 g/day (median 1,928) for pregnant women and 2,242 ± 658 g/day (median 2,164) for lactating women. Tapwater intake was 1,157 ± 635 g/day (median 1,065) for control women, 1,189 ± 699 g/day (median 1,063) for pregnant women, and 1,310 ± 591 g/day (median 1,330) for lactating women. Total water intake was equal to or greater than 3,000 g/day among 7 percent of control women, 11 percent of pregnant women, and 13 percent of lactating women. Tapwater intake was equal to or greater than 2,000 g/day among 10 percent of control women, 15 percent of pregnant women, and 8 percent of lactating women.

Conclusions. These results should be useful in estimating amounts of nutrients and toxic substances that women of reproductive age obtain through the water supply. (*Am J Public Health* 1991;81:328–334)

Introduction

Although weight gain during pregnancy and milk output during lactation cause a theoretical increase in the physiological requirement for water,[1] the actual range of water intake among pregnant and lactating women is not known.[2] The relative importance of water as a dietary source of nutrients such as calcium cannot be evaluated in the absence of intake data. Furthermore, secondary exposure of a fetus or nursing child to toxic substances in the water supply makes accurate data on quantities of water consumed by the mother particularly important in establishing criteria for setting allowances of contaminant levels. To make population-based estimates of quantities and sources of tapwater and total water ingested by pregnant and lactating women, we used data from a large national US Department of Agriculture (USDA) dietary survey,[3–5] which provided an opportunity to consider water intake from all sources including solid foods. For purposes of comparison, we have also estimated intakes for similarly aged non-pregnant, non-lactating women who participated in the same survey.

Methods

Survey Design

In 1977–78 the USDA conducted the Nationwide Food Consumption Survey (NFCS) to gather dietary information on individuals living in randomly sampled nonmilitary households in the contiguous 48 states.[3] The NFCS first sampled households and asked the head of household to complete a seven-day household food use and purchase record. Of the 20,812 households contacted, 14,930 (72 percent) agreed to participate in this activity. Ninety-four percent of the individuals sampled from the participating households completed an interview that included a 24-hour recall; they were then asked to keep a food diary for the next two days.[6]

Participating in the Individual Intake Survey were 30,770 persons.[3] Of these, 87 women stated that they were nursing a child aged two years or less, and 214 women stated that they were four or more months pregnant (women less than four months pregnant were classified by the USDA as not pregnant).[3,4] The pregnant and lactating women interviewed for the NFCS were 15–49 years of age; thus, for potential control subjects, we identified all female NFCS participants 15–49 years old who were neither pregnant nor lactating (n = 7,227).

Analytic Sample

Our data analysis was confined to those NFCS participants having complete and usable data on demographic variables of interest, self-reported weight and height, and all foods, beverages, and drinking water consumed for three consecutive days. Overall, 86 percent of interviewed participants,[7] including 188 pregnant women, 77 lactating women, and 6,201 non-pregnant, non-lactating control women, fit these criteria.

Address reprint requests to Abby G. Ershow, ScD, Nutritionist, Lipid Metabolism-Atherogenesis Branch, National Heart, Lung and Blood Institute, Bethesda, MD 20892. Ms. Brown is an Epidemiologist with the Biostatistics Branch, National Cancer Institute; Dr. Cantor is an Epidemiologist, Environmental Epidemiology Branch, NCI. This paper, submitted to the Journal June 23, 1989, was revised and accepted for publication August 6, 1990.

To compensate for deletions, we developed a weighting factor ("NCI weighting factor") equal to the inverse of the retained (i.e., not deleted) proportion of participants in each age-sex-race-region-season stratum.[7] This NCI weighting factor was derived for the entire data set and applies to the pregnant and lactating participants as well; it did not reflect any assumptions regarding diet or water intake of non-respondents. The retained participants did not differ from the original sample with regard to the distribution of values for age, sex, race, region, season of interview, urbanization, and education of heads of household,[7] and for this reason we consider our analytic sample to be satisfactorily representative of all surveyed participants. All calculations incorporated the NCI weighting factor as well as a USDA weighting factor that reflected the interview completion rate in each sampling stratum.[3,4,7]

Definition of Terms

"Drinking water" refers to plain water consumed directly as a beverage, but does not include tapwater used to prepare foods and other beverages. "Tapwater" includes drinking water as well as tapwater added in final preparation of foods and tapwater-based beverages. "Total water" includes all water from tapwater and non-tapwater sources, including water contained in food.

Calculation of Tapwater and Total Water Intake

The methods, coding procedures, and specific assumptions used in calculating tapwater and total water intake from the 4,545 foods and beverages reported as consumed by NFCS study participants have been fully described elsewhere[7] and will be summarized here. For all food items listed on the surveyed individuals' dietary records,[5,8] an experienced nutritionist (author AGE) estimated the proportion of the total water content likely to derive from household tapwater added during preparation. Examples of this estimated tapwater proportion are: 100 percent (tea, coffee); 84 percent (reconstituted frozen fruit juices); 50 percent (most reconstituted canned soups); 0 percent (meats, fresh whole milk, fresh fruit, carbonated beverages). These tapwater proportion values were merged with data on the water content of foods[5,7] and on the quantities of foods (including soups) and beverages consumed,[4] providing an estimate of total water and tapwater intake from all foods and beverages. To these estimates we added the cups of drinking water consumed each day, yielding the quantities of total water and tapwater from each dietary source for each individual over three days.[7] The original USDA dietary data were recorded in common household units and converted to grams during data processing.[8] Our results are expressed in grams, in keeping with the format of the public use data tapes.[4]

Statistical Methods

All calculations and statistical procedures were conducted using the SAS software package,[9] and incorporated weighting factors (see above) reflecting the interview completion rate in each sample stratum and the deletion of participants with incomplete data.[7] Values represent three-day averages for each individual. All values reported in the text represent mean ± standard deviation. Values for the fraction of total water intake provided by tapwater (see Table 3, below) represent the mean proportion for individuals.[10] Values for the fraction of total water and tapwater intake provided by various dietary sources (see "Results," below) represent the population proportion.[10]

For the entire group of study subjects (n = 6,466), sources of between-individuals variation in daily total water and tapwater intake were examined using analysis of variance (ANOVA) for unbalanced designs (PROC GLM).[9] The model included main-effects terms for reproductive status, race, urban/rural residence, region, and season; and covariate terms for age, weight, height, and mean household educational level. Education of the heads of household was a surrogate measure for income, which many participants (21 percent) would not report. Comparisons of means were made using a multiple comparison test (LSMEANS PDIFF) based on the analysis-of-variance model, in which the significance of each contrast was determined after adjusting for all other main effects and covariates.[9]

Results

Characteristics of Participants

Demographic and physical characteristics of the study participants are shown in Table 1. The women were fairly evenly distributed among the northeastern, southern, midwestern and western regions of the country. Interviews took place during all four seasons of the year. Most of the women were white and lived in urban or suburban rather than rural areas. The lactating women were generally older than the pregnant women and tended to live in households with a higher level of education. Self-reported heights and weights of our study sample compared favorably with physical measurements of height and weight made on adult females participating in the second US National Health and Nutrition Examination Survey (NHANES II).[11]

Total Water Intake

Effect of Reproductive Status. Total water intake of study participants is shown in Table 2. Mean and median daily consumption of total water for all participants was approximately 2 liters (i.e., 2,000 g). The upper ranges of intake, however, were much higher. Seven percent of control women, 11 percent of pregnant women, and 13 percent of lactating women consumed 3 liters (3,000 g) or more total water per day. Three percent, 4 percent, and 3 percent of control, pregnant, and lactating women, respectively, had total daily water intakes equal to or exceeding 3.5 liters (3,500 g).

Reproductive status was significantly associated with total water intake ($p < .001$). Lactating women consumed the largest quantity (g/day) of total water ($p < .001$ compared with control mean), but the mean intake of pregnant women was also significantly elevated ($p < .05$ compared with control women). The total water intake of lactating women was considerably higher than that of pregnant women, but this difference was not statistically significant due to relatively small numbers of participants.

Lactating women also had the highest mean intake of total water when expressed in terms of body weight ($p < .01$ compared with both control and pregnant women). Total water intake per unit of body weight did not differ between pregnant and control women.

Of the 77 lactating women, 64 had infants who also were NFCS study subjects; no information was available regarding the other 13 infants. Total water intake of these 64 lactating mothers was not significantly correlated with either the age (mean = 5 mo) or weight (mean = 7 kg) of their infants.

Total water intake per unit of energy (i.e., g/kcal/day) did not vary as a function of reproductive status. This reflects the parallel increase of water and energy intake: control, pregnant, and lactating women consumed 1,592 ± 561, 1,719 ± 556, and 1,887 ± 626 kcal/day, respectively (mean ± SD).

TABLE 1—Characteristics of Study Participants

	Pregnant Women	Lactating Women	Non-pregnant Non-lactating Women (Controls)
Number of Observations	188	77	6201
Age (yr) (mean ± SD)	25.5 ± 5.6	27.1 ± 4.8	30.4 ± 10.0
Weight* (kg) (mean ± SD)	65.8 ± 10.6	61.9 ± 12.4	61.8 ± 13.0
Height* (cm) (mean ± SD)	164 ± 6	164 ± 6	163 ± 7
Race (%)†			
White	79.6	94.8	81.9
Black	14.5	4.2	13.7
Other	6.0	1.0	4.5
Education** (%)			
Less than high school	20.1	7.9	23.6
High school graduate	37.9	32.3	42.5
More than high school	41.4	59.8	33.4
Urbanization of residence (%)			
Rural	22.4	25.9	26.6
Urban/Suburban	77.6	74.1	73.4
Region of residence (%)			
Northeast	20.8	20.7	25.1
Midwest	29.4	35.6	25.7
South	26.4	13.7	30.3
West	23.3	30.0	18.8
Season interviewed (%)			
Spring	24.6	27.5	22.6
Summer	26.7	26.1	25.7
Autumn	23.1	18.8	26.6
Winter	25.6	27.6	25.1

*By self-report

†Percentages may not add to 100.0%.

**Refers to the female head of household, who usually was the study participant.

Effects of other factors. The effects of factors other than reproductive status on total water intake were also examined in the entire sample of 6,466 women. The most striking effects were due to regional variations in residence. Total water intakes were lowest in the northeast (mean ± SD 1,827 ± 819 g/day; n = 1,551), intermediate in the south (1,955 ± 844 g/day; n = 2,147) and midwest (1,972 ± 912 g/day; n = 1,666), and highest in the west (2,063 ± 1,096 g/day; n = 1,102). Regarding another aspect of residence, rural women consumed more total water than urban/suburban women (1,997 ± 866 g/day, n = 1,885 vs 1,931 ± 925 g/day, n = 4,581).

Effects of race on total water intake may have been due to variation in food and beverage preferences. The small number of non-White, non-Black women had higher total water intakes (2,017 ± 957 g/day, n = 241) than White women (1,953 ± 898 g/day, n = 5,379), who in turn had higher total water intakes than Black women (1,893 ± 960 g/day, n = 846). Seasonal variation in total water intake was due primarily to elevated intake during the summer (2,011 ± 1,006 g/day; n = 1,384). Intakes were very similar in the autumn (1,916 ± 930 g/day; n = 1,617), winter (1,938 ± 960 g/day; n = 1,523), and spring (1,926 ± 762 g/day; n = 1,942).

Tapwater Intake

Effect of reproductive status. Mean and median daily consumption of tapwater for all groups (Table 3) was approximately 1.1 liters (1,100 g). Two liters (2,000 g) or more of tapwater were consumed daily by 10 percent, 15 percent, and 8 percent of control, pregnant, and lactating women respectively. Four percent of control and pregnant women and 3 percent of lactating women ingested 2.5 liters (2,500 g) or more of tapwater daily. Similar to the situation for total water, lactating women had the highest mean daily tapwater consumption ($p < .01$ compared with control women). The difference in tapwater intake between lactating women and pregnant women was nearly as large, but not significant due to small numbers. Pregnant women consumed only slightly more tapwater than control women. When compared with control women, mean tapwater intake per unit of body weight (i.e., g/kg/day) was also significantly higher in lactating women ($p < .05$) but not in pregnant women. Pregnant women actually had a slightly lower intake of tapwater per kg body weight than control women, possibly reflecting weight gain in pregnancy or a shift to sources of fluid (such as milk) that contain relatively little tapwater.

Effects of other factors. Effects on tapwater intake of factors other than reproductive status were similar in direction to those for total water, but the magnitude of the effects was smaller. Region again was the strongest residential factor. Intake was highest in the south (1,228 ± 798) and west (1,208 ± 982) and intermediate in the midwest (1,166 ± 857), whereas women in the northeast consumed less tapwater (1,034 ± 743 g/day) than in all other regions. Rural women consumed more tapwater (1,237 ± 819 g/day) than urban/suburban women (1,132 ± 847 g/day). Tapwater intake of non-White, non-Black women (1,232 ± 815 g/day) was higher than that of either White women (1,159 ± 841) or Black women (1,143 ± 849 g/day). Seasonal intake, which varied only slightly, was highest in the summer (1,198 ± 918), followed by winter (1,165 ± 895), autumn (1,144 ± 861), and spring (1,130 ± 711).

Sources of Total Water and Tapwater

The contribution of various dietary sources to total water intake is shown in Table 4. The reader should note that the mass of the non-water fraction of the foods does not enter into the calculation. For example, pregnant women had a mean intake of total water from milk and milk drinks of 308 g/day (Table 4), but their mean intake of milk *per se* would be a slightly higher figure. The presence of zeros in the 50th percentile column indicates that at least half of the women in that particular group consumed essentially no legumes, nuts or seeds, fruit juice, tea, coffee, soft drinks, or alcoholic beverages over the three days of the study.

Considerable individual variation is demonstrated by the marked differences in intake between the mean, median, and 95th percentiles. For example, the average coffee and tea intake of our subjects was not especially high. Table 4 shows that mean intake for pregnant women was 132 g water/day in tea and 197 g water/day in coffee, whereas lactating women consumed, on average, 253 g water/day in tea and 205 g water/day in coffee (an 8-oz cup = 240 g water). The even lower 50th percentile values indicate that many of the women did not consume any tea or coffee at all. Nevertheless, a small fraction of subjects, representing the top 5 percent of

TABLE 2—Total Water Intake of Women 15–49 Years Old

Reproductive Status*	Mean	Standard Deviation	Percentile Distribution 5	10	25	50	75	90	95
g/day									
Control	1940	686	995	1172	1467	1835	2305	2831	3186
Pregnant	2076	743	1085	1236	1553	1928	2444	3028	3475
Lactating	2242	658	1185	1434	1833	2164	2658	3169	3353
g/kg/day									
Control	32.3	12.3	15.8	18.5	23.8	30.5	38.7	48.4	55.4
Pregnant	32.1	11.8	16.4	17.8	22.8	30.5	40.4	48.9	53.5
Lactating	37.0	11.6	19.6	21.8	28.4	35.1	45.0	53.7	59.2
g/kcal/day									
Control	1.3	0.9	0.7	0.8	0.9	1.2	1.5	2.0	2.5
Pregnant	1.3	0.5	0.7	0.8	0.9	1.2	1.5	1.9	2.2
Lactating	1.3	0.4	0.7	0.8	1.0	1.2	1.5	1.8	2.1

*Number of observations: non-pregnant, non-lactating controls (n = 6,201); pregnant (n = 188); lactating (n = 77).

TABLE 3—Tapwater Intake of Women 15–49 Years Old

Reproductive Status*	Mean	Standard Deviation	Percentile Distribution 5	10	25	50	75	90	95
g/day									
Control	1157	635	310	453	709	1065	1503	1983	2310
Pregnant	1189	699	274	419	713	1063	1501	2191	2424
Lactating	1310	591	430	612	855	1330	1693	1945	2191
g/kg/day									
Control	19.1	10.8	5.2	7.5	11.7	17.3	24.4	33.1	39.1
Pregnant	18.3	10.4	4.9	5.9	10.7	16.4	23.8	34.5	39.6
Lactating	21.4	9.8	7.4	9.8	14.8	20.5	26.8	35.1	37.4
Fraction of daily total water intake (%)									
Control	57.2	18.0	24.6	32.2	45.9	59.0	70.7	79.0	83.2
Pregnant	54.1	18.2	21.2	27.9	42.9	54.8	67.6	76.6	83.2
Lactating	57.0	15.8	27.4	38.0	49.5	58.1	65.9	76.4	80.5

*Number of observations: non-pregnant, non-lactating controls (n = 6,201); pregnant (n = 188); lactating (n = 77).

intake (i.e., the 95th percentile), drank nearly one liter per day of caffeinated beverage.

Sources of tapwater are shown in Table 5. Nearly all the tapwater consumed was provided by a small number of dietary items: drinking water; coffee; tea; grain and grain products; and fruit juices. (Carbonated beverages, most alcoholic beverages, and many fresh foods do not contain tapwater in the first place.) In contrast, a larger range of foods and beverages served as substantial sources of total water (Table 4).

Reproductive status had little effect on the relative contribution (expressed as the population proportion[10]) of drinking water, beverages, and foods to intake of total water (Table 4). For control, pregnant, and lactating women, respectively, the contribution to total water intake was 30 percent, 34 percent, and 30 percent for drinking water; 46 percent, 43 percent, and 45 percent for other beverages; and 24 percent, 23 percent, and 25 percent for foods.

Sources of tapwater intake varied more with reproductive status (Table 5). Drinking water contributed a larger fraction of the tapwater intake of pregnant women (58 percent) than for either control women (50 percent) or lactating women (52 percent). Other beverages accounted for a population proportion of 34 percent, 43 percent, and 41 percent of the tapwater intake of pregnant, control, and lactating women, respectively. Only 7 percent of the tapwater intake of our study participants was contributed by solid foods.

Discussion

By providing population-based estimates of total water and tapwater intake by pregnant and lactating women, our study addresses a nearly complete lack of information on this topic. Recent national surveys of tapwater use in Great Britain[12] and Canada[13] did not present data for pregnant or lactating women. Studies conducted by the California State Department of Health on the relationship of pregnancy outcome and water exposure did not evaluate quantities of water consumed from all sources.[14] A clinical study of fluid intake during lactation was very thorough but was conducted in a single locale (Iowa) using 26 self-selected volunteers.[1,15]

Our estimates derive from a relatively large final analytic sample that included 188 pregnant women and 77 lactating women. NHANES II,[16] which also used a stratified random national sampling design, included even fewer pregnant (n = 122) and lactating (n = 41) women than the 1977–78 NFCS. Larger numbers of pregnant and lactating participants would be desirable but cannot

TABLE 4—Total Water (g/Day) Derived from Various Dietary Sources by Women Aged 15–49 Years*

	Control Women			Pregnant Women			Lactating Women		
		Percentile			Percentile			Percentile	
Sources	Mean†	50	95	Mean†	50	95	Mean†	50	95
Drinking water	583	480	1440	695	640	1760	677	560	1600
Milk and milk drinks	162	107	523	308	273	749	306	285	820
Other dairy products	23	8	93	24	9	93	36	27	113
Meats, poultry, fish, eggs	126	114	263	121	104	252	133	117	256
Legumes, nuts, and seeds	13	0	77	18	0	88	15	0	72
Grains and grain products	90	65	257	96	69	246	119	82	387
Citrus and non-citrus fruit juices	57	0	234	69	0	280	64	0	219
Fruits, potatoes, vegetables, tomatoes	198	171	459	212	185	486	245	197	582
Fats, oils, dressings, sugars, sweets	9	3	41	9	3	40	10	6	50
Tea	148	0	630	132	0	617	253	77	848
Coffee and coffee substitutes	291	159	1045	197	0	955	205	80	955
Carbonated soft drinks**	174	110	590	130	73	464	117	57	440
Non-carbonated soft drinks**	38	0	222	46	0	257	38	0	222
Beer	17	0	110	7	0	0	17	0	147
Wine, spirits, liqueurs, mixed drinks	10	0	66	5	0	25	6	0	59
All Sources	1940	NA	NA	2076	NA	NA	2242	NA	NA

*Number of observations: nonpregnant, non-lactating controls (n = 6,201); pregnant (n = 188); lactating (n = 77).

†Individual means may not add to all-sources total due to rounding.

**Includes regular, low-calorie, and no-calorie soft drinks.

NA: Not appropriate to sum the columns for the 50th or 95th percentiles of intake.

TABLE 5—Tapwater (g/Day) Derived from Various Dietary Sources by Women Aged 15–49 Years*

	Control Women			Pregnant Women			Lactating Women		
		Percentile			Percentile			Percentile	
Sources	Mean†	50	95	Mean†	50	95	Mean†	50	95
Drinking water	583	480	1440	695	640	1760	677	560	1600
Milk and milk drinks	5	0	0	16	0	73	15	0	130
Other dairy products**	—	—	—	—	—	—	—	—	—
Meats, poultry, fish, eggs	15	0	70	12	0	65	15	0	70
Legumes, nuts, and seeds	8	0	45	11	0	54	8	0	54
Grains and grain products	42	20	164	46	21	170	61	35	240
Citrus and non-citrus fruit juices	26	0	138	34	0	161	39	0	153
Fruits, potatoes, vegetables, tomatoes	9	0	65	9	0	56	11	0	57
Fats, oils, dressings, sugars, sweets	4	0	33	4	0	24	3	0	37
Tea	148	0	630	132	0	617	253	77	848
Coffee and coffee substitutes	291	159	1045	197	0	955	205	80	955
Carbonated soft drinks**††	—	—	—	—	—	—	—	—	—
Non-carbonated soft drinks††	27	0	188	31	0	203	21	0	148
Beer**	—	—	—	—	—	—	—	—	—
Wine, spirits, liqueurs, mixed drinks	1	0	0	0	0	0	0	0	0
All Sources	1157	NA	NA	1189	NA	NA	1310	NA	NA

*Number of observations: non-pregnant, non-lactating controls (n = 6,201); pregnant (n = 188); lactating (n = 77).

†Individual means may not add to all-sources total due to rounding.

**Foods and beverages in this category are not sources of tap water.

††Includes regular, low-calorie, and no-calorie soft drinks.

NA: Not appropriate to sum the columns for the 50th or 95th percentiles of intake.

readily be obtained without expensive oversampling procedures.

We think it unlikely that our estimates of water intake are affected by serious non-response bias. The overall household response rate for the 1977–78 NFCS was 72 percent, very similar to the 73 percent individual response rate obtained for NHANES II,[11] and nearly all eligible individuals (94 percent) within participating NFCS households agreed to participate in the individual intake surveys as well. Furthermore, dietary data are probably no more susceptible to bias than other types of health information, and NHANES II investigators found no indication of household non-response bias for health-related variables.[11] In addition, there is no evidence that water intakes of persons with complete data differed systematically from intakes of persons lacking complete data. We thus consider our analytic sample to be representative of the surveyed study groups.

How reproducible (i.e., reliable) and accurate (i.e, valid) are our estimates? Few other data are available, but our estimates compared favorably with them. Median intake of at-home drinking water (plain tapwater) by pregnant women interviewed by the California Department of Health was 2–3 cups per day,[14] very similar to our finding that median drinking water intake of pregnant women was 640 g (2.7 cups) per day (Tables 4 and 5). The lactating women in our study had an average total water intake per unit of energy (1.3 g/kcal/day; Table 2) similar to that reported by lactating women in the Iowa study (1.4 g/kcal/day)[2] and in both studies two-thirds of the participants consumed more than the recommended 1 g/kcal.[1] As expected, our larger national sample of lactating women had a wider range of total water intake values (1,087–4,159 g/day), which completely encompassed the range for the 26 Iowa participants (1,920–3,957 g/day).[2] Both our study and the Iowa study found the highest level of total water intake to be approximately 4 liters (4,000 g) per day.

The non-pregnant, non-lactating control women in our study had a mean intake of total water from beverages (1,480 g/day; Table 4) very similar to that of adult British women (1.5–1.6 L/day).[12] Intake of tapwater from beverages (1,081 g/day; Table 5) also was very similar to that of British women (1.0–1.1 L/day).[12] Canadian women, on the other hand, consumed relatively more tapwater-based beverages (1.3–1.6 L/day).[13]

As the comparisons above indicate, our results are generally compatible with other published data despite differences in study population, intake methodology, and sample size. It is possible, however, that our estimates are on the low side. Using three days' worth of data for each participant would presumably allow the under-reporting biases of 24-hour recalls and the over-reporting biases of dietary records to balance out.[6,17] Nevertheless, energy intakes reported by NFCS participants were lower than expected and were very similar to the energy intakes yielded by the single 24-hour recall used in the NHANES II survey.[18] Under-reporting of food intake thus may account for the lower mean total water intake of our lactating study subjects (2,242 g/day, Table 2) compared with that of the Iowa participants (2,860 g/day).[2] Alternatively, the diets of the highly educated Iowa participants may not have been representative of lactating women's diets in general. The Iowa participants were studied as part of an attempt to increase milk production by deliberately increasing fluid intake,[15] and they may have been consuming extra fluid even during the baseline phase of the study.

How timely are the estimates? The data used for this analysis were collected in 1977–78. In the last decade there have been important changes in certain aspects of the food supply, including increased purchase and consumption of meals outside the home and greater intake of carbonated soft drinks.[19] Unfortunately, the USDA's more recent Continuing Survey of Food Intake by Individuals[19] did not collect information on intake of plain water (i.e., drinking water), and data from these surveys therefore cannot be used to derive either tapwater or total water intake.

Our population-based cross-sectional study confirms that pregnancy and lactation are associated with significant increases in water intake, although a prospective study design is needed to evaluate whether water intake increases spontaneously with the onset of pregnancy and lactation. The estimated theoretical increase in water requirement caused by lactation is 0.5–1.0 L/day[1,20] but in our study the mean daily total water intake of lactating women was only 302 g (i.e., 302 ml) higher than the total water intake of control women (Table 2). The non-pregnant, non-lactating control women in our study consumed quantities of total water well above the minimum amount needed to prevent dehydration. This implies that many if not most women may not need to increase their water intake when they become pregnant or begin to lactate, because their pre-pregnancy intakes were already more than adequate. Unnecessary recommendations to consume "lots of fluids" might inspire some women to replace nutritionally valuable solid foods with water or other nutrient-free beverages such as tea, coffee, or diet soda. Dietary counseling might be better directed towards the few women who habitually consume very small amounts of liquid.

Our evaluation of the relative contribution of various dietary sources to total water and tapwater intake has implications for questionnaire design. Studies of water consumption can be conducted more efficiently by limiting questions to the most important sources. We found that drinking water, coffee, and tea alone accounted for a population proportion[10] of 87 percent of tapwater intake by our study participants (calculated as the sum of source-specific means divided by the all-sources mean, Table 5). Adding reconstituted fruit juices and cooked grain products (such as rice, hot cereal, noodles, pasta) to this list brings this figure to 94 percent. A nearly complete assessment of the quantities of tapwater consumed could thus be made using a relatively short questionnaire. A longer list of sources (Table 4) would be needed to assess total water intake by questionnaire.

We found that women living in rural areas and in the western or southern regions of the United States consumed more total water and tapwater than women living in urban/suburban areas in the northeast and midwest. This could reflect differences in climatic temperature and humidity, physical activity, and diet, particularly beverages. Estimates of intake of water-borne elements might be improved by taking these residential factors into account. Season, on the other hand, was responsible for less variation, probably due to climate control and alternation throughout the year of hot and cold drinks. Similar minor effects of season were observed in Canada, where tapwater intake was slightly (5–10 percent) higher in the winter.[13] The Iowa study[2,15] also found that total water intake by lactating women was minimally affected by season.

Our findings may be used to estimate intake of minerals via tapwater. Calcium is of particular interest because of the very high demands imposed by reproduction; the Recommended Dietary Allowance (RDA) for calcium for most adult women is 800 mg/day, whereas for pregnant and lactating women the RDA is 1,200 mg/day.[1] The calcium content of home water supplies averages 50–60 mg/L, although in hard water areas the concentration may reach 145 mg/liter.[21] In the typical situation, pregnant or lactating women ingesting 1.2–1.3 liters/day of tapwater (Table 3) from the home supply would take in by this route only 60–80 mg/day of calcium, or about 5 percent of the 1,200 mg recommended daily allowance. Only the women living in areas with very hard water (145 mg/L) and drinking unusually large quantities of this hard water (2.2–2.4 L/day; 95th percentile of intake; Table 3) would obtain an appreciable amount of calcium from the tap.

Estimates of exposure to toxic metals and chemicals in tapwater often are made with the assumption that few persons consume more than 2 liters per day[21–24]. Canadian researchers found that 12 percent of adults consumed 2 liters or more of tapwater daily.[13] In our study 10 percent of the non-pregnant, non-lactating women, 8 percent of the lactating women, and 15 percent of the pregnant women consumed 2,000 g/day or more of tap water (see "Results"), and 3–4 percent consumed 2,500 g or more. This suggests that 2 liters/day might not be a sufficiently high estimate of intake to protect 95 percent of these susceptible population groups, particularly pregnant women, from water-borne toxic substances unless a sufficient margin of safety is incorporated into estimates of exposure. ☐

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Regression-Adjusted Small Area Estimates of Functional Dependency in the Noninstitutionalized American Population Age 65 and Over

Jennifer M. Elston, MHSA, Gary G. Koch, PhD, and William G. Weissert, PhD

ABSTRACT

Health planning efforts for the population age 65 and over have been hampered continually by the lack of reliable estimates of the noninstitutionalized long-term care population. Until recently national estimates were virtually nonexistent, and reliable small area estimates remain unavailable. However, with the recent publication of several national surveys and the 1990 Census, synthetic estimates can be made for states and counties by using multivariate methods to model functional dependency at the national level, and then applying the predicted probabilities to corresponding state and county data.

Using the 1984 National Health Interview Survey's Supplement on Aging and the 1986 Area Health Resources File System, we have produced log-linear regression models that include demographic and contextual variables as predictors of functional dependency among the noninstitutionalized population age 65 and over. Age, sex, race, and the percent of the 65 and over population who reside in poverty were found to be significant predictors of functional dependency. Applying these models to 1986 Medicare Enrollment Statistics, regression-adjusted synthetic estimates of two levels of functional dependency were produced for all states and—as examples of how the rates can be used to produce additional synthetic estimates—the largest county in each state. We also produced point estimates and standard errors for the national prevalence of functional dependency among the noninstitutionalized population age 65 and over. (*Am J Public Health* 1990;81:335–343)

Introduction

Bivariate correlates of functional dependency suggested in previous work include age,[1–11] being female,[3,4,6–10,12] being of a race other than White,[3,10] being unmarried or residing with family members,[6,7,10] living alone,[13] having a low income,[7,10,11] and being at the low end of the social class continuum.[7] However, Feller[1] found no significant difference in rates of dependency by gender; Dawson, Hendershot, and Fulton[5] found gender differences to disappear when age structure was taken into account; and Jette and Branch[13] found that among those who lived with others, level of income was inversely related to increasing disability.

A few researchers have also investigated the relationship of functional dependency to other factors with the use of multivariate methods. Among these age,[10,14,15–17] physical performance,[10,14] number of chronic conditions,[14,15] sex,[14] race,[14] emotional performance,[14] health status,[14] prior functional abilities,[10] social class,[15] income,[15] and education[16] have been shown to be significantly related to level of functional dependency.

Yet, health planning efforts for the population age 65 and over continue to be hampered by the lack of reliable data and methods for making population-based estimates at subnational levels. Unger and Weissert's[17] research, which used 1977 data to produce synthetic estimates of the functionally dependent noninstitutionalized age 65 and over population, was forced to rely upon a data set with much more limited measures of dependency than those available on more recent national surveys.

Although several methods exist to produce synthetic estimates, none has been found to be uniformly superior. One well suited method uses a fitted regression model to predict outcomes from correlates of the trait or characteristic of interest. This approach has been widely used.[17–29]

This paper presents log-linear regression models as well as regression-adjusted synthetic estimates of the noninstitutionalized dependent age 65 and over population in each state and its largest county. With appropriate caution the rates generated by the models can be applied to the number of noninstitutionalized age 65 and over individuals in various age, sex, race, and poverty subgroups living in other counties to produce synthetic estimates of the prevalence of dependency for those counties.

Methods

Data Sources and Definitions

Data were drawn from the 1984 National Health Interview Survey's Supplement on Aging, the 1986 Area Resource File System, *State, Regional, and Na-*

Address reprint requests to William G. Weissert, PhD, Professor, Health Services Management and Policy, and Research Scientist, Institute of Gerontology, University of Michigan, School of Public Health, 1420 Washington Heights, Ann Arbor, MI 48109-2029. Ms. Elston is a doctoral student in Health Services Organization and Policy in the School of Public Health at the University of Michigan; Dr. Koch is with the Department of Biostatistics, School of Public Health, University of North Carolina at Chapel Hill. This paper, submitted to the Journal April 14, 1989, was revised and accepted for publication April 9, 1990.

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tional Monthly and Seasonal Heating Degree Days,[30] and 1986 Medicare Enrollment Statistics.

The 1984 Supplement on Aging, based on a multistage area probability sample, provides self-reported characteristics (such as conditions and impairments, living arrangement, and functional abilities) for 11,497 civilian noninstitutionalized people age 65 and over. A more detailed description of the survey is provided by Fitti and Kovar.[31]

Contextual variables were drawn from two sources: the 1986 Area Resources File System and *State, Regional, and National Monthly and Seasonal Heating Degree Days*. The 1986 Area Resource File System is a county-specific data base containing more than 7,000 variables concerning a wide range of social, economic, and health resource characteristics, and is described in detail elsewhere.[32] The use of one additional contextual variable, heating degree days (which is available from the National Oceanic and Atmospheric Administration) was examined.

Using the geographic identifiers available on the 1984 Supplement on Aging, corresponding contextual variables from the above two sources were attached at the Standard Metropolitan Statistical Area (SMSA) for individuals residing in one of 31 large self-representing SMSAs. Individuals on the data set who resided outside these 31 areas were assigned the corresponding regional (northeast, north central, south, or west) and urbanity (SMSA or nonSMSA) average for their type of residence. The result was 39 distinct geographic areas: 31 self-representing SMSAs, and four urban and four nonurban regional areas.

Synthetic estimates are produced by multiplying the regression-adjusted national rates by population data. Because any explanatory variable included in the national model must also be available in the small area population data, Medicare Enrollment Statistics were used for small area population data. Enrollment data by age, sex, and race for each state and county in the United States are available from the Bureau of Data Management and Strategy at the Health Care Financing Administration.[33]

Model Specification

Unit of Analysis

The unit of analysis for this study was the individual participant in the 1984 Supplement on Aging. Although the weighted sample size of the 1984 Supplement on Aging is over 26 million, we normalized the provided survey weight variable to sum to the actual sample size of 11,497 so as not to exaggerate significance levels in model evaluation.

Dependent Variable

The dependent variable for our analysis was a three level hierarchical measure which classified individuals into their highest level of dependency defined as follows.

ADL Dependent: Individuals residing in the community who, because of a health or physical problem, reported that at the time of the survey they had difficulty with and received human assistance with eating, transferring, toileting, dressing, or bathing.

Mobility/IADL Dependent: Individuals residing in the community who, at the time of the survey, were not ADL dependent, but because of a health or physical problem reported difficulty with and received human assistance with inside mobility, outside mobility, meal preparation, grocery shopping, money management, housework (light and heavy), or telephone usage.

Independent: Individuals residing in the community who, at the time of the survey, were neither ADL nor IADL dependent.

Explanatory Variables

Coupling the constraints of the Supplement on Aging and Medicare county data, the following explanatory variables were available for use:

- Sex (coded 1 if female and 0 if male);
- Race (coded 1 if other and 0 if White);
- Age Group (coded as a zero-centered variable equal to the midpoint age in five-year intervals from 65 to 85 and over minus 77.5, divided by 5, i.e. −2, −1, 0, 1, or 2);
- Age-Squared (coded as the square of the "age group" variable, i.e. 4, 1, 0, 1 or 4); and
- Pairwise Interactions (coded as the product of the pair).

In addition, a number of contextual variables were hypothesized to affect the rate of functional dependency among the noninstitutionalized population age 65 and over. These included:

- the number of nursing home beds per 1,000 population age 65 and over;
- the number of unoccupied nursing home beds per 1,000 people age 65 and over;
- the number of acute care hospital beds per 1,000 people age 65 and over;
- the per capita income of the population;
- the percent of the persons age 65 and over who reside in poverty;
- the number of primary care physicians per 1,000 persons age 65 and over;
- the percent of the poverty population that is covered by Medicaid;
- the age-adjusted mortality rate;
- the population per square mile;
- the population age 65 and over per square mile;
- the percent of the population that resides in an urban area; and
- heating degree days (the annual number of degrees of heating or cooling which would be required to move an area's temperature one degree for one day toward 68 degrees Fahrenheit).

Each of the contextual variables was evaluated for entry into the models both continuously and categorically. For the categorical analysis the variables were collapsed into three levels: high, medium, and low—coded −1, 0, and 1, respectively.

Statistical Methods

Given the multinomial nature of the dependent variable, a multicategory extension of logistic regression which provides a general structure was used. (The structure imposed on the data by an ordered logistic regression model was evaluated and found to be inappropriate.) The log-linear model was fit using a SAS supported procedure designed for categorical data modeling, PROC CATMOD.[34] For log-linear model analysis, CATMOD uses maximum likelihood estimation. Given the three category dependent variable, two sets of parameter estimates were produced: one for the logged ratio of not dependent to ADL dependent, and one for the logged ratio of IADL dependent to ADL dependent. Working with these two equations simultaneously yielded a formula for each category of the dependent variable: 1) not dependent; 2) IADL dependent; and 3) ADL dependent.

Because failure to account for the complex sampling design of surveys such as the Supplement on Aging can lead to underestimated variances of the regression coefficients, the magnitude of the design effects was tested with RTILOGIT.[35] Results confirmed previous findings that design effects on the Supplement on Aging

TABLE 1—Regression Results: Predictors of Functional Dependency among the American Noninstitutionalized Age 65 and Over Population[a]

	Log (Prob. not dependent/ Prob. ADL dependent)		Log (Prob. IADL dependent/ Prob. ADL dependent)		
Variables	Coefficient	Standard Error	Coefficient	Standard Error	Chi-Square[b] (d.f. = 2)
Race	−.41	.12	−.02	.13	28.56
Sex	−.17	.08	.71	.09	217.76
Age group	−.59	.03	−.20	.03	654.65
Age group-squared	−.11	.02	−.07	.02	29.23
Poverty	−.35	.07	−.15	.06	36.66
Intercept	2.51	.08	.44	.09	1920.81

[a] Lack of fit chi-square = 128.13, d.f. = 108, p = .0906. Model chi-square = 959.14, d.f. = 10, p = .001.
[b] $p < .001$

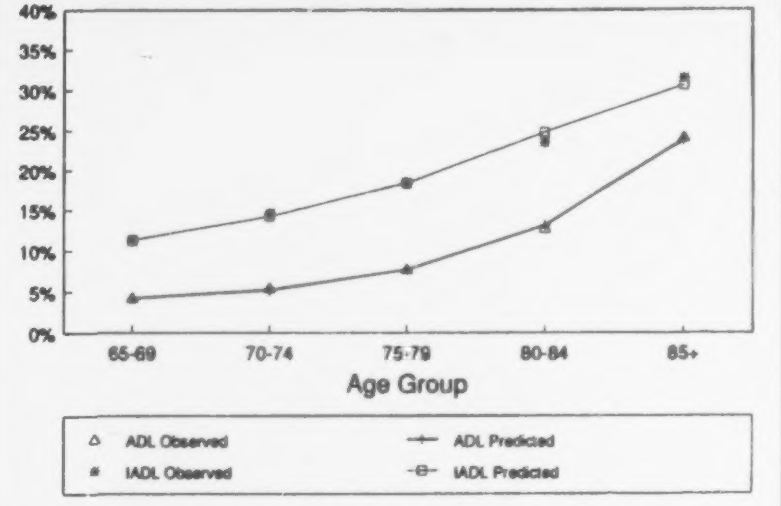

FIGURE 1—Observed and Predicted ADL and IADL Dependency Rates

are relatively small,[31] and estimates produced by complex sample methods such as RTILOGIT would not alter results or conclusions.

For model testing, the data base was randomly divided in half within each primary sampling unit. In the first half of the data base candidate models were fit for the dependent variable. Once model development was completed, the goodness-of-fit of the model was validated in the other half of the data base as described in the Appendix.

Results

Direct Estimates

Direct estimates from the 1984 Supplement on Aging show that in 1984 approximately 2.0 million (7.3 percent) of the noninstitutionalized population age 65 and over in the United States suffered from at least one ADL dependency, and an additional 4.2 million (16.4 percent) suffered from at least one IADL dependency. National prevalence and percentage estimates by race, sex, and age and their accompanying standard errors are presented in Appendix Tables A1-A4.

Regression-Adjusted Results

Table 1 presents the survey-weighted results of the log-linear regression analysis. Five variables were selected to form the final model: race, sex, age, age-squared, and one contextual variable, the categorical variable reflecting the percent of the population age 65 and over who reside in poverty. Each was a significant predictor of functional dependency in the overall model.

Other contextual variables were eliminated from the model. Three of these variables (when entered as the sole contextual variable in the model) had p-values $< .02$ (the number of heating degree days, the ratio of Medicaid recipients to the population below poverty, and the number of unoccupied nursing home beds per 1,000 elderly). However, all p-values became nonsignificant ($p > .05$) when entered in the model in combination with the poverty variable.

While theoretical considerations would argue for inclusion of these and the other contextual variables in the model, inclusion complicates the calculation of synthetic estimates. Since the goal of our study was to provide a practical and easily administered tool for producing synthetic estimates we elected to omit nonsignificant variables from our model.

Similarly, although the continuous form of the poverty variable might seem preferable from a statistical perspective, differences in conclusions between the continuous and the categorical form of the variable were negligible, and as results and examples could more easily be presented with the categorical variable, results presented are for the latter.

The fit of this model was evaluated with the log-likelihood ratio chi-square statistic. Since the statistic was nonsignificant, the use of the model was supported. The need for pairwise interactions of the variables was evaluated and determined to be unnecessary.

Further evaluation of the fit of the model was done as described in the Appendix, and by plotting the observed age-specific rates of dependency and the regression-predicted rates of dependency. Both ADL and IADL predicted rates closely approximate the observed rates (Figure 1). However, when the population

TABLE 2—Percentages of ADL Dependent Americans Age 65 and Over Living in the Community by Age, Sex, and Race: Regression-adjusted Estimates

Race	Sex	Ages 65–69	Ages 70–74	Ages 75–79	Ages 80–84	Ages 85 & Over	Ages 65 & Over
Low (<8%) Poverty Community							
White	Male	2.5	3.2	4.9	9.1	18.8	4.3
	Female	2.8	3.4	5.1	9.0	17.4	5.6
	Both	2.7	3.4	5.1	9.0	17.7	5.1
All Others	Male	3.7	4.6	7.0	12.4	24.1	6.5
	Female	3.9	4.8	7.0	11.7	21.1	6.1
	Both	3.8	4.8	7.0	11.8	23.0	6.2
All races	Male	2.7	3.3	5.1	9.3	19.9	4.5
	Female	2.9	3.6	5.3	9.3	17.5	5.6
	Both	2.8	3.5	5.2	9.3	18.1	5.2
Moderate (8–15%) Poverty Community							
White	Male	3.5	4.4	6.7	12.1	23.9	6.4
	Female	3.8	4.7	6.9	11.7	21.5	7.2
All Others	Male	5.1	6.3	9.4	16.2	29.7	7.9
	Female	5.3	6.4	9.1	14.8	25.5	9.0
	Both	5.2	6.4	9.3	15.3	26.4	8.5
All Races	Male	3.7	4.6	6.9	12.4	24.0	6.5
	Female	3.9	4.8	7.0	11.9	21.7	7.3
	Both	3.8	4.7	6.9	12.0	22.4	7.0
High (>15%) Poverty Community							
White	Male	4.9	6.1	9.1	15.9	29.7	8.4
	Female	5.2	6.3	9.1	14.9	26.2	9.0
	Both	5.1	6.2	9.1	15.3	27.3	8.8
All Others	Male	7.0	8.5	12.5	20.7	35.9	11.4
	Female	7.1	8.4	11.8	18.4	30.2	11.3
	Both	7.1	8.5	12.0	19.0	32.3	11.3
All Races	Male	5.2	6.4	9.7	16.4	30.6	8.9
	Female	5.5	6.7	9.5	15.5	26.7	9.4
	Both	5.3	6.6	9.6	15.8	28.0	9.2

is divided into smaller subgroups, such as females of a race other than White, the model fits somewhat less well, but this could be due to the greater variability in those cases from reduced sample size.

Table 2 presents regression-adjusted estimates of the prevalence of ADL dependency, and Table 3 of IADL dependency. Because the poverty variable has three values—less than 8 percent (low); between 8 and 15 percent (moderate); and over 15 percent (high) of the population age 65 and over residing in poverty—three sets of estimates were produced.

Results showed the relative likelihood of both ADL and IADL dependency increases quadratically with age, is higher if of a race other than White, and increases with an increasing percent of the population age 65 and over residing in poverty. The likelihood of IADL dependence is higher if female, but the relative likelihood of being ADL dependent if female varies across age, race, and poverty groups. Although it is in general higher for females than males until age 80 in communities of low and moderate levels of poverty, and until age 75 for those in high poverty communities, after these ages the percent of noninstitutionalized males with an ADL impairment is either equal to or greater than that of females.

Regression-Adjusted Synthetic Estimates

Although Medicare Enrollment Statistics were the best available population data, two adjustments had to be made. First, because approximately 5 percent of the population age 65 and over is not enrolled in Medicare, the numbers were inflated to reflect the total population age 65 and over. The percent of the population age 65 and over enrolled in Medicare varies little across sex or family income groups, but does differ across race groups,[36] therefore the number of White Medicare enrollees was inflated by 4.4 percent, and the number of all other enrollees was inflated by 13.5 percent.

Second, because Medicare Enrollment data include both the noninstitutionalized and institutionalized age 65 and over population, and rates produced from our analysis are applicable for the noninstitutionalized population only, the numbers therefore were deflated to remove the noninstitutionalized age 65 and over population.

Using the 1985 National Nursing Home Survey and the 1985 National Health Interview Survey, a logistic regression equation was produced to estimate rates of institutionalization among the age 65 and over population at the national level. Candidate explanatory variables for inclusion in the model included those available on the merged data set for which corresponding population data existed. Given this constraint, age (in five-year intervals from 65 to 85 and over), sex, race (White and all others), and geographic region (northeast, north central, south, and west), as well as their pairwise interactions and transformations were available for use. Region was included because the supply of nursing home beds varies geographically.[37] The model found to best fit the data included age, age-squared, sex, and an indicator variable reflecting whether or not the individual resided in the north central region of the country. Estimates produced from the model were used to deflate the state and county population data to be representative of the noninstitutionalized age 65 and over population (results are available from the authors).

By applying the regression-adjusted rates of dependency to the adjusted Medicare data, synthetic estimates of the functionally dependent noninstitutionalized age 65 and over population were produced for each state, and the largest county in each state (Tables 4 and 5).

These estimates are based upon three assumptions. First, that the race, sex, age, and poverty-specific disability rates from the 1984 Supplement on Aging did not change between 1984 and 1986. Second, that the relationship between dependency and race, sex, age, and the percent of the population age 65 and over residing in poverty is the same for each state and its largest county as it is for the nation. And third, that race, sex, age, and the percent of the population age 65 and over residing in poverty are the only important predictors of functional dependency. Thus, the estimates will err to the extent that the relationship between dependency and race, sex, age, and poverty in a community has changed over time; the extent that the relationships vary from national averages; and the extent to which other known or unknown factors

TABLE 3—Percentages of IADL Dependent Americans Age 65 and Over Living in the Community by Age, Sex, and Race: Regression-adjusted Estimates

Race	Sex	Ages 65–69	Ages 70–74	Ages 75–79	Ages 80–84	Ages 85 & Over	Ages 65 & Over
Low (<8%) Poverty Community							
White	Male	5.2	6.6	8.9	12.5	17.3	7.3
	Female	11.7	14.4	18.8	25.2	32.5	17.5
	Both	8.7	11.5	14.5	22.0	29.4	13.5
All Others	Male	7.5	9.3	12.3	16.7	21.7	10.3
	Female	16.2	19.7	25.0	32.0	38.7	21.1
	Both	12.0	17.0	17.7	29.3	27.7	16.7
All Races	Male	5.5	6.8	9.2	12.7	18.2	7.6
	Female	12.2	15.2	19.2	25.9	32.8	17.9
	Both	9.1	12.2	14.7	22.7	29.2	13.8
Moderate (8–15%) Poverty Community							
White	Male	6.3	7.8	10.4	14.4	18.9	9.0
	Female	13.8	16.9	21.7	28.2	34.8	19.9
	Both	10.4	13.1	17.3	23.5	29.9	15.5
All Others	Male	8.9	11.0	14.3	18.8	23.1	11.8
	Female	18.9	22.7	28.2	35.0	40.4	25.2
	Both	13.8	17.6	21.4	28.9	36.6	19.1
All Races	Male	6.5	8.1	10.7	14.6	19.1	9.2
	Female	14.1	17.3	22.0	28.6	35.1	20.2
	Both	10.6	13.4	17.5	23.8	30.2	15.7
High (>15%) Poverty Community							
White	Male	7.5	9.3	12.2	16.3	20.3	10.5
	Female	16.2	19.6	24.7	31.1	36.5	22.1
	Both	12.4	15.2	19.7	25.9	31.1	17.3
All Others	Male	10.5	12.8	16.3	20.8	24.1	14.1
	Female	21.7	25.8	31.4	37.6	41.3	28.3
	Both	17.3	21.1	25.5	33.5	35.0	23.1
All Races	Male	7.9	9.8	12.9	16.8	20.9	11.0
	Female	17.0	20.8	25.9	32.3	37.1	23.2
	Both	13.1	16.2	20.7	27.1	31.6	18.3

which are not in the model strongly influence functional dependency.

Discussion

Using log-linear models, regression-adjusted synthetic estimates of the rates of two levels of functional dependency among the noninstitutionalized age 65 and over population were produced. Age, sex, race, and the percent of the population age 65 and over who reside in poverty were found to be significant predictors of ADL and IADL dependency among the noninstitutionalized population age 65 and over.

Introduction of a contextual variable into the regression model appears overdue given that estimated rates are substantially affected by the poverty variable. The results here, which are consistent with other researchers' work, suggest that just as poverty is a strong correlate of many unwanted problems in youth and adulthood, so too its sequelae are present in old age, manifesting as higher dependency rates.

Results here also confirm the strong relationship reported by other researchers between dependency and age, as well as the variation in age-specific rates of dependency between men and women, and Whites and all other races. While explication of the underlying determinants of these variations is beyond the scope of this paper, reconfirming their importance suggests the need for policies and research agendas sensitive to these relationships and variations. Of particular importance is the quadratic relationship between age and dependency, meaning that with each passing five-year interval rates of dependency increase at an increasing rate—a sobering prospect given the rapid expansion of the oldest old population.

The percentages generated with the regression models (Tables 2 and 3) can be multiplied by corresponding population estimates for specific geographic areas of interest to generate regression-adjusted synthetic estimates of dependency (similar to those in Tables 4 and 5) among the noninstitutionalized age 65 and over population.

Such estimates are likely to be most useful as a gauge of the size of the long-term care population in subnational areas. They may also be of use as initial building blocks for estimating demand for formal long-term care services. While functional dependency estimates will not translate directly to service demand without other important measures (i.e. income, price, preferences, and other need measures such as social support), previous research has shown that utilization of formal health care services is closely related to need.[38,39] Furthermore, the usefulness of the methods outlined here will be greatly enhanced by the 1990 US Census, which will make age, sex, race, and poverty data widely available.

Finally, it should be noted that while the analysis supports the use of these equations to produce small area estimates of functional dependency among the noninstitutionalized age 65 and over population, the quality of the small area estimates produced by them still needs to be evaluated in future research using geographic indicators (not presently available for public use) which permit comparison of synthetically produced estimates with direct estimates from adequately large samples for the same small areas. □

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TABLE 4—1986 Dependent Noninstitutionalized Population Age 65 and over by State: Regression-adjusted Synthetic Estimates

	Percent Dependent			Number Dependent		
State	Total	ADL	IADL	Total	ADL	IADL
District of Columbia	33.3	11.4	21.9	23,805	8,134	15,671
Mississippi	30.1	10.4	19.7	89,953	31,064	58,889
Louisiana	29.3	10.1	19.2	125,130	42,983	82,147
Alabama	28.9	9.8	19.1	136,220	46,280	89,940
Georgia	28.6	9.6	19.0	163,071	54,646	108,425
South Carolina	28.2	9.4	18.8	95,562	31,845	63,717
Missouri	28.2	9.7	18.5	180,495	61,902	118,593
Arkansas	28.1	9.7	18.5	90,470	31,139	59,331
Oklahoma	28.1	9.7	18.5	107,760	37,049	70,711
Tennessee	28.1	9.5	18.5	154,822	52,591	102,231
South Dakota	28.0	9.8	18.2	26,148	9,186	16,962
Virginia	28.0	9.5	18.6	158,247	53,509	104,738
North Carolina	28.0	9.4	18.6	195,934	65,655	130,279
Texas	27.8	9.5	18.3	409,903	139,822	270,081
Kentucky	27.7	9.5	18.2	115,895	39,612	76,283
Maine	27.6	9.5	18.1	41,482	14,291	27,191
North Dakota	27.5	9.7	17.9	22,777	7,993	14,784
West Virginia	27.3	9.3	18.0	65,472	22,317	43,155
New Mexico	26.7	9.2	17.5	36,127	12,405	23,722
Idaho	26.3	9.1	17.2	27,933	9,641	18,292
Hawaii	25.8	8.3	17.5	27,006	8,674	18,332
Nebraska	24.0	7.7	16.4	47,939	15,291	32,648
New York	24.0	7.5	16.6	521,526	162,389	359,137
Kansas	23.9	7.6	16.3	72,273	22,868	49,405
Iowa	23.7	7.5	16.2	91,976	29,141	62,835
Minnesota	23.6	7.5	16.1	114,774	36,478	78,296
Massachusetts	23.6	7.3	16.3	177,057	55,050	122,007
Maryland	23.5	7.2	16.3	103,331	31,767	71,564
Illinois	23.5	7.3	16.2	300,132	92,840	207,292
Rhode Island	23.3	7.2	16.1	31,443	9,707	21,736
Vermont	23.2	7.3	15.9	14,244	4,470	9,774
Indiana	23.1	7.1	15.9	141,704	43,800	97,904
Delaware	23.0	7.1	15.9	15,969	4,920	11,049
Connecticut	23.0	7.1	15.9	92,857	28,771	64,086
Wisconsin	23.0	7.2	15.8	136,679	42,885	93,794
New Jersey	23.0	7.1	15.9	215,410	66,269	149,141
Ohio	23.0	7.1	15.9	284,862	87,738	197,124
Pennsylvania	23.0	7.1	16.0	376,538	115,873	260,665
New Hampshire	22.9	7.1	15.8	26,237	8,155	18,082
Colorado	22.9	7.2	15.7	63,173	19,751	43,422
Michigan	22.9	7.0	15.8	229,029	70,624	158,405
Florida	22.7	7.1	15.6	426,861	133,050	293,811
Wyoming	22.6	7.1	15.5	8,928	2,814	6,114
Oregon	22.6	7.1	15.5	77,330	24,273	53,057
Washington	22.6	7.1	15.5	111,303	34,826	76,477
Montana	22.5	7.1	15.4	20,813	6,567	14,246
Utah	22.3	7.0	15.3	27,495	8,577	18,918
Arizona	21.8	6.7	15.1	82,697	25,501	57,196
Alaska	21.5	6.6	14.8	3,676	1,138	2,538
Nevada	20.4	6.2	14.2	19,249	5,845	13,404
California	19.4	5.5	13.9	521,891	147,505	374,386

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TABLE 5—1986 Dependent Noninstitutionalized Population Age 65 and over by County: Regression-adjusted Synthetic Estimates

County	Percent Dependent			Number Dependent		
	Total	ADL	IADL	Total	ADL	IADL
New Orleans, LA	31.5	10.8	20.7	19,335	6,630	12,705
Fulton, GA	31.1	10.5	20.6	18,991	6,416	12,575
Hinds, MS	30.6	10.5	20.2	8,150	2,787	5,363
Shelby, TN	30.4	10.4	20.0	23,878	8,138	15,740
Jefferson, AL	30.2	10.3	19.9	25,716	8,756	16,960
Baltimore City, MD	30.1	10.2	19.9	31,189	10,547	20,642
Philadelphia, PA	29.7	10.0	19.7	71,559	24,182	47,377
Pulaski, AR	29.0	9.8	19.1	10,363	3,520	6,843
Dade, FL	28.9	10.0	18.9	66,997	23,245	43,752
Oklahoma, OK	28.2	9.6	18.6	18,351	6,231	12,120
Greenville, SC	27.4	9.1	18.3	9,490	3,144	6,346
Marion, IN	27.3	7.8	19.6	16,774	4,757	12,017
Honolulu, HI	25.9	8.2	17.6	19,656	6,263	13,393
Denver, CO	24.5	7.7	16.8	15,395	4,838	10,557
Queens, NY	24.4	7.5	16.8	61,693	19,097	42,596
Wayne, MI	24.2	7.5	16.7	59,915	18,469	41,446
Douglas, NE	24.1	7.5	16.6	9,993	3,109	6,884
Jefferson, KY	24.1	7.4	16.7	19,164	5,889	13,275
Multnomah, OR	24.1	7.6	16.5	19,274	6,069	13,205
Hennepin, MN	24.1	7.6	16.6	25,701	8,071	17,630
Mecklenburg, NC	24.0	7.3	16.7	10,008	3,041	6,967
Los Angeles, CA	24.0	7.5	16.5	184,822	57,601	127,221
Cass, ND	23.8	7.6	16.2	2,056	655	1,401
Cook, IL	23.8	7.3	16.5	135,009	41,486	93,523
Polk, IA	23.7	7.3	16.4	7,964	2,464	5,500
Harris, TX	23.7	7.2	16.4	38,936	11,922	27,014
Providence, RI	23.6	7.3	16.3	20,369	6,300	14,069
Chittenden, VT	23.5	7.3	16.2	2,276	704	1,572
Minnehaha, SD	23.5	7.4	16.1	3,007	947	2,060
Cumberland, ME	23.5	7.3	16.2	6,954	2,171	4,783
Cuyahoga, OH	23.4	7.2	16.3	47.719	14,604	33,115
New Castle, DE	23.2	7.1	16.0	10,037	3,083	6,954
King, WA	23.2	7.2	15.9	34,343	10,715	23,628
Sedgwick, KS	23.1	7.1	16.0	9,250	2,848	6,402
New Haven, CT	23.1	7.2	16.0	24,590	7,607	16,983
Laramie, WY	23.0	7.3	15.7	1,303	411	892
Hillsborough, NH	23.0	7.1	15.9	7,336	2,259	5,077
Kanawha, WV	22.8	7.0	15.8	6,925	2,121	4,804
Salt Lake City, UT	22.5	7.0	15.5	11,630	3,613	8,017
Yellowstone, MT	22.3	7.0	15.4	2,552	794	1,758
Henrico, VA	22.3	6.7	15.6	2,813	848	1,965
Ada, ID	22.3	6.9	15.4	4,059	1,261	2,798
Bernalillo, NM	22.2	6.8	15.3	9,109	2,800	6,309
Maricopa, AZ	22.0	6.8	15.2	46,918	14,419	32,499
Anchorage, AK	20.2	6.0	14.2	1,168	346	822
Clark, NV	20.2	6.1	14.1	11,017	3,321	7,696
Middlesex, MA	20.1	5.7	14.4	32,144	9,062	23,082
Milwaukee, WI	19.9	5.6	14.3	24,335	6,812	17,523
St. Louis, MO	19.4	5.4	14.0	21,853	6,088	15,765
Bergen, NJ	18.9	5.3	13.6	21,774	6,080	15,694

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APPENDIX

To assess goodness-of-fit the model was first applied to the second half of the data set, and its goodness-of-fit was evaluated with the chi-square statistics associated with the individual parameters and with the lack-of-fit statistic.

Second, the model was applied to the entire sample to test the fit of the estimated coefficients. This was done by including an indicator variable representing the half of the data set from which each observation came, as well as all of its pairwise interactions with the other variables in the model.

Third, goodness-of-fit of the model was evaluated by comparing the similarity of the model-predicted dependency rates with their observed counterparts in the other half of the data set. In so doing, the candidate models were used to determine the predicted values of the probability of dependence for individuals in the other half of the data base. The differences between these predicted values and their true value gives a residual value for that individual. The closeness of the averages of the residuals to zero for various subgroups of individuals (e.g. males, females, different age groups, etc.) and lack of correlation of the residuals with characteristics of individuals are indicative of goodness-of-fit. In addition, Pearson correlations were evaluated for the residuals and each of the explanatory variables, and their low values supported the fit of the model.

TABLE A-1—Direct Point Estimates of the Number of Noninstitutionalized Americans Ages 65 and over Who Were Functionally Dependent in 1984 by Age, Sex, and Race

		Personal Care Dependent[a]						Mobility or Household Activity Dependent[b]					
Race	Sex	Ages 65–69	Ages 70–74	Ages 75–79	Ages 80–84	Ages 85 & Over	Ages 65 & Over	Ages 65–69	Ages 70–74	Ages 75–79	Ages 80–84	Ages 85 & Over	Ages 65 & Over
White	Male	164,644	143,520	147,240	87,750	111,319	654,473	255,589	238,009	185,456	121,775	114,223	915,052
	Female	157,176	190,273	192,626	234,942	286,274	1,061,291	608,833	629,901	668,386	479,112	422,742	2,808,976
	Both	321,820	333,793	339,866	322,692	397,593	1,715,764	864,422	867,910	853,844	600,887	536,965	3,724,028
Other	Male	20,095	15,920	27,803	17,060	12,628	93,506	34,700	35,507	12,834	12,037	6,461	101,539
	Female	27,466	29,173	32,021	24,656	30,263	143,579	88,035	129,402	93,806	60,226	33,796	405,265
	Both	47,561	45,093	59,824	41,716	42,891	237,085	122,735	164,909	106,640	72,263	40,257	506,804
All Races	Male	184,739	159,440	175,043	104,810	123,947	747,979	290,289	273,516	198,290	133,812	120,684	1,016,591
	Female	184,642	219,446	224,647	259,596	316,537	1,204,870	696,868	759,303	762,194	539,338	456,538	3,214,241
	Both	369,381	378,886	399,690	364,408	440,484	1,952,849	987,157	1,032,819	960,484	673,150	577,220	4,230,832

[a] Personal care dependent includes bathing, dressing, toileting, transferring, or eating. Individuals classified as personal care dependent may also be dependent in mobility or household activities but are counted only as personal care dependent.

[b] Mobility or household activity includes inside mobility, outside mobility, meal preparation, grocery shopping, money management, housework (heavy and light), and telephone usage. Individuals already classified and counted in this table as personal care dependent are excluded from this category.

SOURCE: 1984 National Health Interview Survey's Supplement on Aging.

TABLE A-2—Direct Point Estimates of the Percent of Noninstitutionalized Americans Ages 65 and over Who Were Functionally Dependent in 1984 by Age, Sex, and Race

		Personal Care Dependent[a]						Mobility or Household Activity Dependent[b]					
Race	Sex	Ages 65–69	Ages 70–74	Ages 75–79	Ages 80–84	Ages 85 & Over	Ages 65 & Over	Ages 65–69	Ages 70–74	Ages 75–79	Ages 80–84	Ages 85 & Over	Ages 65 & Over
White	Male	4.45	5.29	7.70	9.68	20.49	6.70	7.1	9.0	9.9	13.8	21.6	9.6
	Female	3.46	5.08	6.63	13.44	23.55	7.50	13.7	17.1	23.5	27.9	36.1	20.3
	Both	3.91	5.17	7.05	12.16	22.61	7.17	10.7	13.7	18.1	23.1	31.6	15.9
Other	Male	5.29	5.63	12.40	20.11	30.64	9.23	9.5	13.0	6.4	15.4	15.7	10.4
	Female	5.87	6.39	11.51	13.00	31.38	9.65	19.7	29.5	34.3	32.3	38.1	28.3
	Both	5.61	6.10	11.91	15.20	31.16	9.48	15.1	23.2	21.9	27.3	31.0	21.1
All Races	Male	4.53	5.32	8.19	10.58	21.20	6.93	7.3	9.3	9.5	13.9	21.2	9.7
	Female	3.69	5.22	7.06	13.40	24.13	7.70	14.2	18.4	24.5	28.3	36.2	21.0
	Both	4.06	5.26	7.51	12.44	23.23	7.34	11.1	14.6	18.5	23.5	31.5	16.4

[a] Personal care dependent includes bathing, dressing, toileting, transferring, or eating. Individuals classified as personal care dependent may also be dependent in mobility or household activities but are counted only as personal care dependent.

[b] Mobility or household activity includes inside and outside mobility, meal preparation, grocery shopping, money management, housework and laundry, or taking medications. Individuals already classified and counted in this table as personal care dependent are excluded from this category.

SOURCE: 1984 National Health Interview Survey's Supplement on Aging.

(*continued*)

APPENDIX (cont.)

TABLE A-3—Standard Errors for Direct Point Estimates of the Number of Noninstitutionalized Americans Ages 65 and over Who Were Functionally Dependent in 1984 by Age, Sex, and Race

		Personal Care Dependent[a]						Mobility or Household Activity Dependent[b]					
Race	Sex	Ages 65–69	Ages 70–74	Ages 75–79	Ages 80–84	Ages 85 & Over	Ages 65 & Over	Ages 65–69	Ages 70–74	Ages 75–79	Ages 80–84	Ages 85 & Over	Ages 65 & Over
White	Male	22,089	17,716	19,053	16,788	15,575	44,178	28,512	35,599	19,739	16,863	15,080	51,398
	Female	61,455	20,109	23,692	21,521	25,062	24,852	41,039	34,866	52,187	28,654	31,635	97,799
	Both	30,234	32,062	31,102	27,462	30,537	81,762	52,840	44,439	60,723	30,779	37,923	114,393
Other	Male	6,615	6,695	9,789	6,726	6,028	15,914	9,982	11,011	6,021	5,462	4,759	18,232
	Female	19,032	8,025	8,729	9,101	7,743	8,847	14,699	18,944	16,384	13,098	9,883	38,966
	Both	9,418	11,479	12,331	10,894	11,348	26,045	19,473	24,630	16,333	15,095	10,408	50,672
All Races	Male	23,550	19,188	22,393	18,136	16,076	47,292	29,841	26,794	21,340	18,231	15,813	53,851
	Female	21,067	23,577	25,280	26,080	26,048	62,634	42,659	37,894	55,004	31.530	31,851	102,984
	Both	33,080	32,690	36,685	28,888	31,919	86,074	56,564	49,549	63,438	35,726	38,704	124,479

[a] Personal care dependent includes bathing, dressing, toileting, transferring, or eating. Individuals classified as personal care dependent may also be dependent in mobility or household activities but are counted only as personal care dependent.

[b] Mobility or household activity includes inside mobility, outside mobility, meal preparation, grocery shopping, money management, housework (heavy and light), and telephone usage. Individuals already classified and counted in this table as personal care dependent are excluded from this category.

NOTE: Standard errors calculated using SAS supported procedure PROC SESUDAAN.[40]

SOURCE: 1984 National Health Interview Survey's Supplement on Aging.

TABLE A-4—Standard Errors for Direct Point Estimates of the Number of Noninstitutionalized Americans Ages 65 and over Who Were Functionally Dependent in 1984 by Age, Sex, and Race

		Personal Care Dependent[a]						Mobility or Household Activity Dependent[b]					
Race	Sex	Ages 65–69	Ages 70–74	Ages 75–79	Ages 80–84	Ages 85 & Over	Ages 65 & Over	Ages 65–69	Ages 70–74	Ages 75–79	Ages 80–84	Ages 85 & Over	Ages 65 & Over
White	Male	.60	.61	.94	1.97	2.73	.45	.71	.91	.10	.19	2.70	.47
	Female	.45	.64	.69	1.23	1.68	.40	.84	.91	1.59	1.67	2.01	.62
	Both	.37	.48	.59	.98	1.48	.31	.58	.65	1.16	1.23	1.75	.41
Other	Male	1.80	2.32	4.22	6.84	11.67	1.43	2.50	3.75	2.64	6.09	9.27	1.62
	Female	1.81	2.02	3.26	3.78	8.08	1.16	2.68	3.04	4.27	5.12	8.24	1.90
	Both	1.23	1.61	2.23	3.48	7.06	.91	2.09	2.69	2.70	4.32	6.48	1.52
All Races	Male	.58	.60	.98	1.96	2.66	.43	.67	.87	.97	1.79	2.62	.45
	Female	.44	.55	.74	1.17	1.60	.37	.82	.85	1.51	1.49	1.85	.58
	Both	.37	.44	.62	.94	1.43	.30	.57	.63	1.08	1.13	1.61	.40

[a] Personal care dependent includes bathing, dressing, toileting, transferring, or eating. Individuals classified as personal care dependent may also be dependent in mobility or household activities but are counted only as personal care dependent.

[b] Mobility or household activity includes inside and outside mobility, meal preparation, grocery shopping, money management, housework and laundry, or taking medications. Individuals already classified and counted in this table as personal care dependent are excluded from this category.

NOTE: Standard errors calculated using SAS supported procedure PROC SESUDAAN.[40]

SOURCE: 1984 National Interview Survey's Supplement on Aging.

Dioxins and Dibenzofurans in Adipose Tissue of US Vietnam Veterans and Controls

Han K. Kang, DrPH, Kevin K. Watanabe, MS, Joseph Breen, PhD, Janet Remmers, MS, Margaret G. Conomos, MPH, John Stanley, PhD, and Michele Flicker, PhD, MD

ABSTRACT

The primary reason for concern about the adverse effects of exposure to Agent Orange is attributable to its toxic contaminant, 2,3,7,8-tetrachlorodibenzo-p-dioxin (TCDD) or dioxin. We studied adipose tissues from 36 Vietnam veterans, a similar group of 79 non-Vietnam veterans, and 80 civilians; the tissue specimens were selected from the 8,000 archived tissues collected from the non-institutionalized general population by the US Environmental Protection Agency. The geometric mean (± standard deviation) dioxin levels in adipose tissue for Vietnam veterans, non-Vietnam veterans, and civilian controls were 11.7 (±1.7), 10.9 (±1.7), and 12.4 (±1.9) parts per trillion on a lipid weight basis, respectively. The mean levels for these groups were not significantly different from each other with or without adjustment for age of individuals, body mass index, and specimen collection year. In addition, none of the surrogate measures of Agent Orange exposure such as military branch, service within specific geographic region, military occupation, and troop location in relation to recorded Agent Orange spray was associated with the dioxin levels in adipose tissue of Vietnam veterans. Our results suggest that heavy exposure to Agent Orange or dioxin for most US troops was unlikely. (*Am J Public Health* 1991; 81:344–349)

Introduction

From 1965 to 1970, the US Air Force sprayed more than 40 million liters of Agent Orange in South Vietnam.[1] Approximately two million American soldiers served in Vietnam during this period. Agent Orange was the name used for a phenoxy herbicide consisting of a mixture of 2,4-dichloro-phenoxyacetic acid (2,4-D) and 2,4,5-trichlorophenoxyacetic acid (2,4,5-T). The 2,4,5-T contained 1-50 parts per million of the contaminant 2,3,7,8-tetrachlorodibenzo-p-dioxin (TCDD), also known as dioxin. Dioxin is extremely toxic in laboratory animals, promotes liver tumors in rats, and is teratogenic in mice.[1,2] The two parent herbicides, 2,4-D and 2,4,5-T are very short-lived in the human body and do not persist in the environment.[3] Because of these characteristics and their relatively low toxicity in humans, attention has been focused on the highly persistent and toxic chemical, dioxin.

Many Vietnam veterans believe that they were heavily exposed to Agent Orange and that the exposure is responsible for health problems such as skin rashes, rare types of cancer, and birth defects of their children. Since 1979, 200,000 Vietnam veterans have come to Department of Veterans Affairs (VA) hospitals for an Agent Orange Registry medical examination because of concerns about exposure to Agent Orange.

Two questions are paramount in dealing with the veterans' concerns. The first is whether dioxin causes birth defects, immune deficiencies, cancers, or other chronic health problems in humans. The second is whether Vietnam veterans, in general, received substantial exposure to dioxin while they served in Vietnam. This study addresses the second question.

Dioxin accumulates preferentially in the body fat of animals and man. Its half-life in humans is estimated at five to 11 years.[4-7] Among the many 2,3,7,8-substituted polychlorinated dibenzo-p-dioxins (PCDDs) and polychlorinated dibenzofurans (PCDFs) found in the environment and in samples of human adipose tissue, only 2,3,7,8-TCDD or dioxin was present in Agent Orange.[7] Dioxin levels in adipose tissues could, therefore, serve as a biological marker of exposure to Agent Orange. In fact, several studies have reported that even after approximately 20 years, dioxin levels were elevated markedly in the adipose tissue or blood serum of Vietnam veterans who handled Agent Orange.[6-8] More recently, however, the Centers for Disease Control (CDC) reported that current serum dioxin levels of Army Vietnam combat troops did not differ significantly from those of non-Vietnam veterans and that dioxin levels in Vietnam veterans did not increase with increased exposure levels estimated from military records.[9] This study was criti-

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Address reprint requests to Han K. Kang, DrPH, Department of Veterans Affairs, Office of Environmental Epidemiology, 1825 K Street, NW, Suite 322, Washington, DC 20006. Mr. Watanabe is also with the Department of Veterans Affairs; Dr. Breen, Ms. Remmers, and Ms. Conomos are with the US Environmental Protection Agency, Washington, DC; Dr. Stanley is with Midwest Research Institute, Kansas City, MO; Dr. Flicker is with the VA Medical Center, Leavenworth, KS. This paper, submitted to the Journal January 16, 1990, was revised and accepted for publication June 19, 1990

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cized on the basis that blood specimens taken almost 20 years after exposure could not represent what happened in Vietnam and that veterans selected for the study were limited to Army enlisted men who served in only one geographic region of Vietnam.[10]

The purpose of our study was twofold: 1) to determine if a group of individuals with military service in Vietnam have significantly higher levels of dioxin in adipose tissue than either a similar group of non-Vietnam veterans or civilian peers, and 2) to determine if dioxin levels in adipose tissue were associated with specific demographic and military service characteristics.

Methods

Selection of Study Subjects

Our study used adipose tissue specimens that had been collected from the general population by the US Environmental Protection Agency (EPA). The EPA has conducted the National Human Adipose Tissue Survey (NHATS) since 1970 to monitor the human body burden of pesticides and other selected chemicals. Up to 1,000 adipose tissue specimens have been collected annually from pathologists and medical examiners across the country and analyzed by the EPA for the selected chemicals. After analysis, the unused tissue specimens were sent to a central facility to be stored at 0°C to −20°C.

The NHATS utilized a probability sample of the Standard Metropolitan Statistical Areas (SMSA) that is designated to represent a sample of the US population in terms of age, sex, and race.[11] The target population for the NHATS program was all non-institutionalized persons in the conterminous US. However, due to the invasive nature of collecting adipose tissue samples, the sampling population was limited to individuals who died from external causes (90 percent) and surgical patients (10 percent). Within each SMSA, hospitals or medical examiners were identified and asked to contribute tissue specimens according to the design specifications of age (0–14 years, 15–44 years, 45+ years), sex and race (White, non-White). Since the vast majority of Vietnam veterans were men born between 1936 and 1954, this study was restricted to specimens from men born in that period.

An Inventory File was created for the 8,000 specimens that were recorded to have an adequate amount of tissue. It was found that 528 of 8,000 specimens were from males born between 1936 and 1954. The hospitals or medical examiners who originally collected the 528 specimens were recontacted to obtain enough identifying information on the donors to determine their military service status. The collection effort yielded information for 494 or 94 percent of the 528 specimens.

The military service status for these 494 men, including any Vietnam service, was determined by reviewing records archived at the National Personnel Records Center (NPRC) in St. Louis and military records maintained at other locations. From this effort, 134 men (or 27 percent of 494) were initially found to have served in the military, 40 of whom served in Vietnam. According to the 1980 Census, one would expect about 34 percent of men in this age group to be veterans. All 40 Vietnam veterans were selected for the study. From the 94 remaining veterans, 80 were selected randomly for the non-Vietnam comparison group. Two civilians were closely matched to each Vietnam veteran by birth year (±two years) and sample collection year (±two years). Age and sample collection year were considered important matching variables because of probable accumulation of dioxin in the body with each year of exposure and the possibility of degradation of fat or dioxin while in storage. All adipose tissue specimens were analyzed during 1987. Demographic data were taken from the NHATS file and the official death certificates. Body mass index (BMI) was calculated from weight and height as follows: BMI = weight in Kg/height in m^2.

Determination of Opportunity for Agent Orange Exposure

A precise estimate of the exposure of each Vietnam veteran to Agent Orange is not considered feasible based on either military records or self-reported data. In this study the probable opportunity for exposure was determined from the following: branch of service, military occupational specialty code (MOSC), and location of the individual's unit in Vietnam in relation to recorded Agent Orange spray. Ground troops in Vietnam might have had a higher probability of contact with Agent Orange than other Vietnam veterans due to the nature of their military operations. Ground troops engaged in combat were assumed more likely to be in herbicide-sprayed areas. As another surrogate for exposure, a veteran's military unit was assigned to one of the four broad military regions in Vietnam. According to the records of the US Air Force Ranch Hand Operation, 20 million liters of Agent Orange were sprayed in Military Region III from 1965 to 1970. During the same period, Military Regions II, I, and IV received 9.5, 8.3, and 4.5 million liters, respectively. Finally, troop locations (company size units) were determined on a 100-meter grid map of Vietnam at intervals of 90 days or less. Computer matching of troop location with respect to time and distance from recorded herbicide spray tracts was carried out using the HERBS tape and services HERB tape databases. The HERBS tape contained information on most of the herbicide missions flown by fixed-wing aircraft from 1965 to 1971, and on crop destruction missions flown by helicopter between 1968 and 1971. The tape contained information on the type of herbicide, gallons, dates, and where spray runs started and ended. Services HERBS Tape prepared by US Army and Joint Services Environmental Support Group identified and documented an additional six million liters of herbicide sprayed mainly by Army personnel around the perimeter of base camps, fire bases, air bases, and other fixed military installations. The opportunity for Agent Orange exposure was determined in two ways: an individual's company was ever located either within two kilometers (km) of a recorded Agent Orange spray tract within three days of application or within eight km of a spray tract within 90 days application.

Laboratory Analysis

Specimens from the three groups were assigned randomly to one of 20 batches. Each batch typically consisted of 10 study specimens and four quality control (QC) samples. The QC samples provided data on method accuracy and precision. In addition, the external quality control audit samples were prepared by another laboratory and incorporated as blind samples into the various batches. Study specimens and QC samples were coded with a unique laboratory number and submitted to the analysts as blind samples. Following sample cleanup and preparation, instrumental analyses were achieved using a high resolution gas chromatograph coupled to a high resolution double-focusing mass spectrometer. Overall method accuracy and precision for 2,3,7,8-TCDD analyses of spiked lipid samples were 113 percent recovery and 8.8 percent coefficient of variation.

The analytical protocol provided for the detection and quantitative determination of 17 congeners of 2,3,7,8-substituted

TABLE 1—Distribution of 2,3,7,8-TCDD Levels in Adipose Tissue by Military Service Status, in pg/g of the Total Extractable Lipid (ppt)

Status	N	Arithmetic Mean ± SD[a]	Geometric Mean	Percentile 25th	 50th (median)	 75th	 90th	 95th
Vietnam Veterans	36[b]	13.4 ± 7.4	11.7	7.8	10.0	17.3	26.8	30.4
Non-Vietnam Veterans	79[c]	12.5 ± 7.2	10.9	7.6	11.4	14.8	19.8	25.3
Civilians	80	15.8 ± 14.5[d]	12.4	7.9	11.8	18.0	30.5	43.4

[a] Standard deviation

[b] Four of the 40 men initially classified as having served in Vietnam were excluded from analysis because two veterans' specimens had less than 20% extractable lipid content, one veteran did not have adequate amount of tissue for analysis, and one veteran served only in Thailand.

[c] One of the 80 men initially classified as having served in the military was excluded because his military service could not be documented unequivocally.

[d] The large standard deviation was attributed to an outlier with a value of 106. The value was verified. The occupational history of this individual is unknown. He was listed as a "laborer" on his death certificate. Analysis conducted without this value resulted in an arithmetic mean of 14.7 (± 10.3) and a geometric mean of 12.2. There was still no statistically significant difference between the groups (p = 0.49).

polychlorinated dibenzo-p-dioxins (PCDD) and dibenzofurans (PCDF). The minimum measurable concentration ranged from 1 picogram per gram (pg/g) for 2,3,7,8-TCDD and 2,3,7,8-TCDF, to 5 pg/g for OCDD and OCDF based on a 10 gram aliquot of human adipose tissue. This protocol was evaluated for method performance prior to being used in this study.[12]

Statistical Considerations

Multiple comparisons and testing for differences were done by using the F test in one way analysis of variance (ANOVA) and analysis of covariance with adjustments for demographic variables such as age, collection year, and body mass index.[13] A paired t-test was conducted to compare the means of Vietnam veterans with their matched civilian controls.[14] In all analyses, the dioxin values were transformed to a natural logarithmic scale because the dioxin values were found to have approximately log-normal distributions in this study and another study.[15]

A stepwise linear regression model was also used to determine whether dioxin levels were associated with demographic and military service characteristics.[16] Factors considered *a priori* as covariates were age, adipose tissue sample collection year, race, and body mass index. A regression model specific to Vietnam veterans included such covariates as military occupation, calendar year of tour in Vietnam, geographic region in Vietnam, time of and distance from recorded Agent Orange spray, and adipose tissue sample collection year.

Results

Table 1 presents the arithmetic and geometric means, and various percentile values for dioxin for the three study groups. One-way analysis of variance did not demonstrate differences in the mean dioxin levels among groups ($p = 0.35$). Analysis of covariance, testing the effect of Vietnam service on dioxin levels after adjusting for age, sample collection year, or body mass index did not indicate an association between service in Vietnam and dioxin levels. A paired t-test between Vietnam veterans and their matched civilian pairs did not show a difference in mean dioxin levels (95% confidence interval for the difference between two means = −1.32,1.16).

For Vietnam veterans, dioxin levels were also evaluated by four factors related to the likelihood of Agent Orange exposure (Table 2). None of the surrogate measures of Agent Orange exposure was associated with the dioxin levels in adipose tissue of Vietnam veterans. Furthermore, the mean dioxin level of seven Vietnam veterans whose specimens were taken within four years (less than one dioxin half-life estimated for humans) since their last service in Vietnam were compared to 19 non-Vietnam veterans whose tissue specimen were collected on or before 1974 (the last sample collection year for the seven Vietnam veterans), and also to their matched civilian pairs. The geometric mean dioxin levels (± standard deviation) for the Vietnam group ($n = 7$), non-Vietnam veteran group ($n = 19$), and civilian controls ($n = 14$) were at the levels of 16.6 (±1.6), 15.5 (±1.5), and 18.4 (±1.6) parts per trillion (ppt), respectively. There was no difference among the means ($p = 0.56$). The 95% CI for a mean dioxin difference between Vietnam veterans and civilian controls were −1.59 and 1.31. Stepwise linear regression analysis for 36 Vietnam veterans indicated that Vietnam service characteristics could account for only 14 percent of the variation in dioxin levels ($p = 0.3$), whereas collection year alone could account for 21 percent of variance ($p = 0.005$). Five other 2,3,7,8-substituted dioxins and 10 other dibenzofurans were measured and their mean levels were calculated from specimens with levels above the detection limit (Table 3). There were no group differences in the mean level of any of the dioxin congeners. The differences in TCDD levels in the three groups were also evaluated while adjusting for levels of other 2,3,7,8-substituted dioxin congeners not found in Agent Orange by a stepwise multiple regression technique. TCDD levels were taken as the dependent variable and Vietnam service status and other dioxin congener levels as independent variables. There was a significant association between the TCDD levels and the levels of PCDD congeners ($R^2 = 0.54$, $p = 0.0001$), but not with the Vietnam service status ($R^2 = 0.0025$, $p = 0.3$).

The levels of dioxins increased with an increase in the number of chlorine except for 1,2,3,7,8,9-HxCDD. Levels of dibenzofurans were always lower than their dioxin counterparts. In each study group, the levels of dioxin tended to be inversely related to the specimen collection year, i.e. the earlier the collection year, the higher the dioxin levels ($p = 0.0001$, $n = 195$).

Discussion

In this study, military service in Vietnam was not associated with elevated dioxin levels in adipose tissue with or without adjustment for demographic variables. In addition, no Vietnam service characteristics measured singularly or in combination was a good predictor of dioxin levels in adipose tissue. There were no consistent trends in the dioxin levels ac-

TABLE 2—Geometric Mean 2,3,7,8-TCDD Levels in Adipose Tissue by Vietnam Service Characteristics, in pg/g of the Total Extractable Lipid (ppt)

Service Characteristics	No. of Vietnam Veterans	2,3,7,8-TCDD Mean	SD[a]
Branch			
Army	20	11.6	1.7
Marine	6	12.3	1.8
Air Force	1	6.7	
Navy	9	12.4	1.5
MOSC[b]			
Non-Combat	24	11.1	1.7
Combat	12	13.1	1.6
Military Region[c]			
I Corp	11	12.4	1.6
II Corp	4	6.9	1.3
III Corp	9	11.9	1.9
IV Corp	3	14.3	1.9
Sea Duty	8	13.3	1.4
Time and Distance From[d] Recorded Herbicide Spray			
a. 3 days/2 KM			
no	31	11.5	1.7
yes	4	14.3	1.6
b. 90 days/8 KM			
no	16	11.8	1.5
yes	19	11.8	1.8

[a] Standard deviation

[b] Military Occupation Specialty Code (MOSC)

[c] Missing Military Region information for one veteran

[d] Missing information on time and distance from recorded herbicide spray for one veteran due to unknown unit location

cording to the surrogate measures of Agent Orange exposure. The mean levels of dioxin did rise slightly with combat MOSC and having been within three days/two km of sprayed areas. But the other two surrogates, branch of service and service location in Vietnam, were not associated with the dioxin levels. The small magnitude of the mean difference and the variation within each group suggest that the small difference in mean values could have been easily due to the sampling and measurement variation.

Analyses of adipose tissue from the general population of industrialized countries have indicated the presence of a number of 2,3,7,8-substituted dioxins and dibenzofurans at ppt level.[17-21] These dioxins and dibenzofurans could have originated from a number of sources: incineration of municipal waste and wood products; manufacturing, use and disposal of pesticides, herbicides and wood preservatives; and PCBs from electric transformers and capacitors. Because 2,3,7,8-TCDD was the only congener found in Agent Orange as a contaminant[1,7] knowing the levels of other dioxins and dibenzofurans would help determine whether TCDD levels in adipose tissue of Vietnam veterans were the result of Agent Orange exposure in Vietnam or exposure to other sources. For example, if most dioxins and dibenzofurans as well as 2,3,7,8-TCDD levels were found to be elevated among Vietnam veterans, contribution from sources other than Agent Orange should not be ruled out. However, if only the 2,3,7,8-TCDD level remained elevated and other PCDD levels were comparable to the comparison groups, Agent Orange would be considered as the contributor. There were no group differences in the mean levels of any PCDD congeners. The lack of group differences suggested that sources other than Agent Orange may have contributed to the Vietnam veterans' levels of 2,3,7,8-TCDD. In a study by Kahn, *et al*,[7] of 10 Vietnam veterans with heavy potential for exposure to Agent Orange (e.g. Ranch Hand personnel, Army Chemical Corps specialists), the levels of 12 other 2,3,7,8-substituted dioxins and dibenzofurans were similar to the levels found among 10 Vietnam veteran controls and seven non-Vietnam veteran controls. Only 2,3,7,8-TCDD levels were elevated approximately 10-fold among Vietnam veterans with a heavy exposure potential.

This study may have failed to detect a small difference in mean dioxin levels among groups because of the relatively small sample size. The study had an adequate statistical power (90 percent) to detect a mean difference of 5 ppt or more between groups. Elimination of dioxin from the body after Vietnam service is an unlikely explanation because dioxin levels of seven Vietnam veterans whose specimens were taken within four years from their return from Vietnam (which was considerably less than the estimated half-life of dioxin in humans) were not significantly different from their appropriate comparison groups.

Although the NHATS sampling scheme was designed to collect a representative sample of the SMSA in terms of age, sex and race, subjects selected for the study may not have represented their respective groups for several reasons. First, over 90 percent of the NHATS sample were collected from deceased persons whose death in most instances were due to traumatic injury. Second, tissue samples for this study were selected from the archived NHATS specimens rather than original NHATS samples. Third, 6 percent of the men who were eligible for the study had to be excluded because of missing personal identifiers. Despite these problems, demographic and military characteristics of the Vietnam veterans selected for the study did not differ substantially from the overall Vietnam veteran population. They were predominantly white (75 percent), draft eligible during the Vietnam war (age 18 to 25), enlisted men (89 percent), who served in the Army and Marine Corps (72 percent) with military occupational specialties related to combat support roles (67 percent). None of the Vietnam veterans in the study had a record of routinely handling or spraying Agent Orange.

The mean background levels of dioxin reported here were generally higher than the values reported by others.[6-9] It is unlikely that this difference is due solely to an interlaboratory variation. Our laboratory (Midwest Research Institute) participated in an interlaboratory validation study for dioxin measurement and produced a satisfactory result. Furthermore, external quality control audit samples prepared by the Battelle Columbus Division were incorporated as blind samples into the various batches and analyzed for dioxin and dibenzofuran congeners. The analytical results were found to be acceptable. We believe that the difference in sampling years between this study and several other studies could account for the higher values reported in our study. We found that the dioxin levels in adipose tis-

TABLE 3—Arithmetic Mean Levels of Dioxins and Furans Detected in Adipose Tissue by Military Service Status, in pg/g of the Total Extractable Lipid (ppt)

Chemicals	Status: Vietnam Veterans	Status: Non-Vietnam Veterans	Status: Civilians
Dioxins			
2378-TCDD	13.4 (36)*	12.5 (79)	15.8 (80)
12378-PeCDD	20.6 (36)	18.3 (78)	18.3 (80)
123478/123678-HxCDD	170.4 (36)	152.9 (79)	165.1 (80)
123789-HxCDD	19.4 (35)	17.2 (79)	17.9 (79)
1234678-HpCDD	276.2 (36)	244.6 (79)	300.3 (80)
OCDD	1261.8 (36)	1108.9 (79)	1392.9 (80)
Furans			
2378-TCDF	2.9 (25)	2.4 (52)	3.3 (51)
12378-PeCDF	1.7 (8)	1.1 (17)	1.9 (16)
23478-PeCDF	23.1 (35)	22.2 (78)	23.3 (80)
123478-HxCDF	21.5 (36)	19.3 (78)	23.2 (79)
123678-HxCDF	10.7 (34)	9.9 (77)	12.0 (79)
234678-HxCDF	3.8 (26)	3.2 (73)	3.6 (78)
123789-HxCDF	1.5 (3)	0.9 (2)	0.9 (4)
1234678-HpCDF	37.4 (36)	32.9 (79)	39.1 (80)
1234789-HpCDF	2.2 (14)	1.9 (35)	2.2 (41)
OCDF	3.6 (27)	4.5 (54)	3.4 (60)

*the number in parentheses represent the number of cases in that category

sue were significantly associated with the sample collection year ($p = 0.0001$), the earlier the collection year, the higher the levels of dioxin irrespective of veteran status. In fact, this general trend was observed for other dioxins. The median sample collection year in our study was 1978, whereas in other studies the specimens were mostly collected in mid 1980s or later. The observed dioxin decline from 1971 to 1982 is consistent with the general trend for chlorinated hydrocarbon chemical compounds in human adipose tissue. The US EPA's National Human Adipose Tissue Survey Program indicates that the median levels of BHC, HCB, and PCB had been steadily decreasing over [illegible] between 1970 and 1983.[11] In Sweden the levels of dioxin and dibenzofurans in human milk decreased significantly from 1972 to 1985.[22] The Swedish authors attributed the decline to the reduction in use of certain organochlorine compounds such as PCBs, PCP and 2,4,5-T. A study involving a large sample of specimens representative of the US population will be needed to confirm this observation.

We concluded that the results of our study did not support the hypothesis that most US troops were heavily exposed to dioxin in Vietnam. Furthermore, none of the surrogate measures of Agent Orange exposure based on military service characteristics was associated with the dioxin levels in adipose tissue of Vietnam veterans. These results are consistent with those of CDC[9] and not inconsistent with Kahn, *et al*,[7] and Schecter, *et al.*[8] ☐

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Kellogg Grant to Strengthen Community Health Services

A grant of $1.2 million has been awarded to the Colleges of Nursing and Medicine at the University of Illinois at Chicago (UIC) by the W. K. Kellogg Foundation to strengthen community-based health service. Emphasizing leadership training for project directors, community members and health care professionals, the project will bring together community residents and health care professionals to create partnerships in advancing leadership in community-based primary health care.

Both classroom instruction and practical hands-on experience at community-based primary health care sites will be offered. UIC faculty and community leaders will serve as mentors for program participants.

For more information, contact Susan Poslusny, Office of Leadership for Primary Health Care in Urban Communities, College of Nursing (m/c 802), the University of Illinois at Chicago, 845 S. Damen, Chicago, IL 60612; (312) 413-0068.

The UCLA Population Studies of CORD: X. A Cohort Study of Changes in Respiratory Function Associated with Chronic Exposure to SO_x, NO_x, and Hydrocarbons

Roger Detels, MD, MS, Donald P. Tashkin, MD, James W. Sayre, DrPH, Stanley N. Rokaw, MD, Frank J. Massey, Jr., PhD, Anne H. Coulson, and David H. Wegman, MD

ABSTRACT

Two never-smoking cohorts in Southern California, one in Lancaster (N = 2340) exposed only to moderate levels of oxidants and the other in Long Beach (N = 1326) exposed to high levels of SO_x, NO_2, hydrocarbons and particulates completed spirometry and the single-breath nitrogen test five to six years apart. Forty-seven percent and 45 percent of the participants were retested. Mean results at baseline for those tested and not retested were similar. Loss to follow-up was primarily due to moving (39 percent and 47 percent). Every difference of consequence indicated greater deterioration in lung function in Long Beach. The level of significance of the difference was greatest, even in the youngest age groups, for $\Delta N_{2_{750-1250}}$, suggesting that the earliest site of impairment may occur in the small airways. Greater deterioration in spirometric parameters was observed in every age group in Long Beach females above seven years of age at baseline and in Long Beach males above 15 years of age, suggesting that chronic exposure to the pollutant mix occurring in Long Beach ultimately adversely affects the large airways as well as small airways. (*Am J Public Health* 1991;81:350–359)

Introduction

The relationship of temporary high levels of SO_x and particulates to health has been well documented by episodes such as those reported in the Meuse Valley, Donora, and London. A higher prevalence of symptoms of chronic obstructive respiratory disease and/or impaired lung function has been reported in cross-sectional studies in areas exposed to high levels of SO_x, NO_x, and/or particulates,[1-4] while in other studies a relationship has not been observed.[5] A number of cohort studies have also been completed, especially in adults.[6-14] The major problems encountered in cohort studies have been maintaining continuous measurement of pollutant levels and controlling for confounding factors which may change. These studies have depended primarily on the use of spirometry for measurement of lung function because of the difficulties associated with doing population studies using possibly more sensitive tests such as the single-breath nitrogen test. There is considerable evidence from experimental studies of both animals and humans that SO_2 and particulate sulfates may be responsible for bronchoconstriction,[15-24] possibly mediated through histamine release and subsequent action on bronchial smooth muscle.[25] A number of these studies have shown such effects at concentrations that occur in ambient air in some urban areas in the United States.

Previously, we reported a greater rate of deterioration in spirometric indices of the single-breath nitrogen test in residents chronically exposed to high levels of oxidants, oxides of nitrogen, sulfate, and particulates compared to an area exposed to low levels of these pollutants.[26]

In this paper we compare changes over a five- to six-year period in spirometric tests and the single-breath nitrogen test in never-smoking residents of an area chronically exposed to high levels of SO_2, sulfates, NO_x, and hydrocarbons but low levels of oxidants with changes to those in residents of another area exposed to low levels of these pollutants to which comparisons to the oxidant-polluted community were made in the previous paper.

Methods

Establishment of Cohorts

Study areas were selected in Lancaster (located 75 miles north of downtown Los Angeles) which historically has been chronically exposed to moderate levels of photochemical oxidants and very low levels of other pollutants and in Long

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Address reprint requests to Roger Detels, MD, MPH, Professor of Epidemiology, School of Public Health, University of California, Los Angeles, CA 90024–1772. All authors were affiliated with UCLA at the time the study was done. Dr. Tashkin is Professor of Medicine/Pulmonary Disease, School of Medicine, and Director, Pulmonary Function Laboratory, UCLA; Dr. Sayre is Principal Statistician and Adjunct Assistant Professor of Biostatistics, UCLA School of Public Health; Dr. Rokaw (deceased) was Clinical Professor of Medicine/Pulmonary Disease, UCLA School of Medicine, and Director, Lung Association of Los Angeles County; Dr. Massey is Professor of Biostatistics and of Biomathematics, UCLA Schools of Public Health and Medicine; Ms. Coulson is Senior Lecturer in Public Health and Research Epidemiologist, UCLA School of Public Health; Dr. Wegman is Professor and Department Head, Department of Work Environment, University of Lowell, Lowell, MA.

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TABLE 1—Demographic Characteristics of Study Census Tracts in Lancaster and Long Beach*

Characteristics	Lancaster	Long Beach
Total residents, all ages	7,069	4,992
White (non-Spanish-surnamed) (%)	6,430 (91.0)	4,939 (98.9)
Spanish-surnamed (%)	434 (6.1)	0 (0.0)
Black (%)	91 (1.3)	3 (0.1)
Other (%)	114 (1.6)	50 (1.0)
Total 7+ years of age	6,121	4,691
Median income	$11,631	$11,474
Number of housing units	2,238	2,197
Proportion of homeowners (%)	63	64
Median home value	$18,600	$23,400

*According to the 1970 census.

Beach (located approximately 30 miles southwest of downtown Los Angeles) which has been chronically exposed to high levels of sulfur dioxide, sulfates, oxides of nitrogen, and probably hydrocarbons. Using 1970 census information, study areas were selected that were demographically similar, each of which contained an established air quality monitoring station. A comparison of the demographic characteristics of the two study areas according to the 1970 census is shown in Table 1. Although the Long Beach study area included a lower proportion of non-Whites, all comparisons in this paper are restricted to White, non-Spanish-surnamed residents who never smoked in order to reduce non-comparability between the cohorts. Participants were considered "never smokers" if they had never smoked more than one cigarette per day. Participants who reported smoking at either the baseline or follow-up examination were excluded from the analysis.

Cross-sectional studies conducted in the Lancaster study area in 1973–74 and in the Long Beach study area in 1974–75 demonstrated poorer lung-function in Long Beach.[27] Concurrent testing was not possible due to funding limitations. Pollutant levels in Long Beach, however, remained similar over this period of time. Details of the recruitment of the study areas have been reported previously.[26–30] Briefly, all the heads of households were identified through reverse telephone directories and voter registration files. Before explanatory letters were sent out, descriptive information regarding the program was presented in the local media. Following the mailing of individualized letters, a neighborhood representative set up an appointment to complete a roster of all household members seven years of age and older and to arrange appointments for lung function testing at the Mobile Lung Function Laboratory, which was located within walking distance.

At the Mobile Laboratory, an interview schedule modified from the National Heart, Lung and Blood Institute Questionnaire was administered; this contained questions concerning respiratory symptoms, and history of respiratory disease, occupation, smoking, and residence. Height and weight were recorded, and the following pulmonary function tests were done: body plethysmography; single-breath nitrogen washout curve, including calculation of $\Delta N_{2_{750-1250}}$* and closing volume; and electronic spirometry measurement with permanent recording of the entire maximal expiratory flow volume relationship on computer tape. The test equipment used was the same for all the testing except for the recorder for the spirometer. The new recorder, however, was thoroughly calibrated to the old recorder before being put into use. Details of the instrumentation used and the procedures followed are contained in previous papers.[28,29] Results of body plethysmographic tests were inconsistent at baseline; thus, changes in these tests could not be determined with confidence. Details of the procedures for administering the single-breath nitrogen test and spirometry are given in Appendix A.

Retesting of the Cohorts

Approximately five years in Lancaster and six years in Long Beach after the baseline studies, retesting in each of the study areas was completed (1978–79 in Lancaster and 1980–81 in Long Beach). Results, however, have all been expressed as annualized rates of change. In the interval between baseline testing and retesting, addresses had been updated through periodic mailings to all participants. Study participants were recontacted and invited to be retested in the same month in which they had originally been tested. Individuals who had moved but remained within a reasonable distance were also encouraged to undergo retesting at the Mobile Laboratory. Intensive efforts were made to persuade all participants to be retested, including call back within one day of individuals who did not keep their appointments at the Mobile Lung Function Laboratory. The procedures for retesting were identical to those used at baseline; the tests were administered in the same sequence and in the same manner.

Individuals who had moved out of the original study area were requested to undergo retesting at the Mobile Laboratory if they visited the Southern California area. Participants who had moved outside of the study area to a known address and who could not be retested were asked to complete a brief questionnaire on respiratory symptoms, smoking, and reasons for moving from the study area, including a final question that asked if the reason for moving was related to respiratory problems.

Validation and Quality Control

Procedures were implemented to calibrate and maintain the reliability of the lung function tests and to assess the effects of concurrent levels of air pollutants. These have been described in detail previously.[29,31] Briefly, they included the following:

Calibration: At the beginning of the testing period in each area the results of lung function testing in the Mobile Laboratory were compared with results obtained from the UCLA (University of California at Los Angeles) Pulmonary Function Laboratory on 10 to 20 participants who were tested on the same day in both laboratories in a randomized sequence. At the beginning of each day's testing in the field and at regular intervals during the testing, the test equipment was recalibrated using standard procedures. In addition, calibration values were recorded on computer tape and evaluated at the end of each test day at UCLA. The Mobile Laboratory was immediately notified of any signs of error or "drift" so that if needed, the instruments could be adjusted before testing commenced the next day. Individuals tested the previous day were

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*Change in the nitrogen concentration between 750 and 1250 cc of expired air.

TABLE 2—Average Number of Days on which Specified Pollutants Were Above California Ambient Air Quality Standards and Means of Daily Peak Hourly Values, 1972–81

Pollutant	Lancaster	Long Beach
SO_2		
Days ≥ 0.05 ppm recorded 1978–81*	0	5
mean (ppm)	.01	.04
Oxidants		
Days ≥ 0.08 ppm for 1972–77, ≥ 0.10 ppm for 1978–81	92	18
mean (ppm)	.07	.04
NO_2		
Days ≥ 0.25 ppm	0	23
mean (ppm)	.03	.11
SO_4		
Days (not recorded)	—	—
mean (µg/m^3, 24 hour totals measured every 6th day, 1977–81)	4.3 (4.3)**	11.3 (11.6)**
Total Suspended Particulates (µg/m^3)		
Days ≥ 100 µg/m^3:		
recorded Lancaster 1978–81	18 (19)**	—
recorded Long Beach 1976–81	—	27 (24)**
annual geometric mean (µg/m^3)	85 (76.1)**	101 (98.6)**

*24 hour standard
**median value

asked to undergo retesting; results for those not retested were deleted.

Reliability and Validity: The reliability of the test procedures was evaluated by requesting every tenth individual completing lung function testing to undergo repeat lung function testing within 20 minutes and comparing the results. In addition, a 3 percent probability sample of participants was asked to undergo additional testing at the UCLA Pulmonary Function Laboratory, usually within two to three months of field testing. The results of the tests in the two laboratories were compared.[26,29]

Effect of Concurrent Pollutant Levels: The possible effect of levels of air pollutants on the day of testing was evaluated by comparing lung function test results in 40 participants in Lancaster and 35 participants in Long Beach tested three times at four-month intervals in each study area. In addition, O/P (observed/predicted — adjusted for age, height, and weight) forced expiratory volume in one second (FEV_1) for all of the adult participants retested was analyzed relative to the level of pollutants recorded for the day on which their test was performed.[30]

Levels of Air Pollutants

Monitoring stations of the Southern California Air Quality Monitoring District located in the two study areas continuously recorded levels of total oxidants, nitric oxide, nitrogen dioxide, total oxides of nitrogen, total hydrocarbons (Lancaster only), and sulfur dioxide.[31] In addition, 24-hour totals for total suspended particulates (TSP) and for sulfates (at six-day intervals) were monitored. The maximum hourly level for oxidants, SO_2 and NO_x, and the 24-hour total for TSP and SO_4, as recorded at the respective monitoring station on the day of testing, were attached to each individual's computer record. All measurements met both the California and federal Environmental Protection Agency guidelines for acceptability of testing procedures for these pollutants.

Results

Levels of Pollutants during the Study Period

The annual means of the daily peak hourly values for SO_2, oxidants, NO_2, and of the 24-hour total value for SO_4 and TSP over the 10-year period are shown in Table 2. The table also gives the average number of days per year on which SO_2, oxidants, NO_2, and TSP exceeded the California ambient air quality standard. Days exceeding these levels were not recorded for all years between 1972 and 1981.[31,32] The years contributing to the determination of the mean number of days are given separately for each pollutant. In terms of both days above California ambient air quality standards and means of the daily peak hourly levels, the largest differences in recorded pollutants were in levels of NO_2. Levels of all indicated pollutants except oxidants, however, were higher in Long Beach. On the other hand, there was a five-fold greater number of days on which oxidants exceeded 0.08 or 0.10 ppm in Lancaster, but less than a two-fold difference in the annual means of the daily peak hourly levels between the two communities. Although the level of oxidants in Lancaster was lower than in most other areas of the Southern California air basin, it was nonetheless higher than in most urban areas in the United States. Unfortunately, levels of hydrocarbons were not measured at the Long Beach monitoring station, but the location of the study area directly downwind from the major oil refining area in Los Angeles and the results of interpolation from adjacent monitoring stations suggest that levels of hydrocarbons were also very high in the Long Beach study area.

Response Rates

Response rates for the two cohorts are given in Table 3. The proportion remaining in the study area who completed all of the lung function tests was reasonably high in both areas (75 percent and 77 percent). An additional 6 percent and 9 percent, respectively, completed only the questionnaire. One percent of participants in both areas were known to have died. The major problem in follow-up was the high proportion of individuals who moved out of the study area: 39 percent (910/2340) in Lancaster and 47 percent (629/1326) in Long Beach. In Lancaster, only 12 percent (270/2340) and in Long Beach only 7 percent (94/1326) of participants refused to be retested. The overall retest rate, excluding deaths, for persons completing *all* lung function tests was 47 percent for Lancaster and 45 percent for Long Beach.

The proportions retested, stratified by age, sex, and place of residence, are given in Table 4. The proportion retested tended to be lower for individuals who were 11–24 years of age at baseline testing, a finding not unexpected in this highly mobile age group leaving home for college and jobs over the five- to six-year interval between baseline testing and retesting.

Because the overall retest rate was not as high as we would have liked, a comparison was made in Table 5 of the lung function test results at baseline among those who completed retesting and those who did not. The high mean observed/predicted values at baseline for those retested reflects the fact that they repre-

TABLE 3—Retest Status of Never Smokers 7–59 Years of Age at Baseline

		Retested						Not Retested			
		Lung Function Test		Questionnaire Only		Deaths		Refused		Lost	
Residence at Baseline	Total Tested at Baseline N	N	(%)	N	(%)	N	(%)	N	(%)	N	(%)
Lancaster											
Study area	1430	1061	(74)	80	(6)	19	(1)	270	(19)	—	—
Moved	910	38	(4)	325	(36)			—	—	547	(60)
Total	2340	1099	(47)			19	(1)				
Long Beach											
Study area	697	523	(75)	62	(9)	18	(3)	94	(13)	—	—
Moved	629	63	(10)	136	(21)			—	—	430	(68)
Total	1326	586	(44)			18	(1)				

TABLE 4—Percent of Never Smokers* Retested, Stratified by Age, Sex, and Place of Residence

	Lancaster		Long Beach	
Sex/Age (years)†	Tested at Baseline N	% Retested	Tested at Baseline N	% Retested
Males				
7–10	210	50	87	40
11–14	217	47	67	27
15–18	140	26	50	32
19–24	61	38	63	38
25–59	262	54	238	53
Females				
7–10	162	46	77	44
11–14	207	51	73	36
15–18	158	35	52	39
19–24	82	37	89	42
25–59	482	55	401	48

*White, non-Spanish-surnamed only, 7–59 years of age at baseline; had not changed job or residence because of a respiratory problem; FEV_1 exists at both times.

†Age at baseline.

sent volunteers who survived for five to six years and were still willing to be retested. Thus, the mean of their observed/predicted test performance at baseline does not reflect the mean for the entire group tested at baseline.

Validity of Results

The mean test/retest results for those 10 percent immediately retested were within 5 percent for the major spirometric indices. Regression analysis of the level of air pollutant on day of testing and the mean test value for FEV_1 for that day for all participants revealed no statistically significant correlations. Of the 24 comparisons made of results of FVC, FEV_1, FEF_{25-75}, FEF_{50-75}, $\Delta N_{2_{750-1250}}$, and closing volume in the two groups of participants tested at different seasons of the year, only one, $\Delta N_{2_{750-1250}}$, was significant at the $p < 0.04$ level; equivalent to chance for this number of comparisons. Mean test results in the cohort retested in three different seasons were not significantly different.

An estimate of the possible impact of errors resulting from the testing procedures in the mobile lung function laboratory was calculated by comparing the results of FEV_1 in the 3 percent probability sample of adults tested in both the mobile and the UCLA lung function laboratories. (Levels of pollution at UCLA fall between those in Lancaster and Long Beach except that the levels of oxidants are higher at UCLA than in either study area.) The mean interval between testing in the mobile and UCLA laboratories was several months.

The difference between FEV_1 values at the Mobile Laboratory were greater at T_2 (retesting) than at T_1 (baseline) in both communities although mobile laboratory values were lower at both examinations. If the UCLA laboratory results represented the true values, then the greater inter-laboratory differences at T_2 than at T_1 suggest that the annual decline in FEV_1 as measured in the field laboratory might have been exaggerated. Furthermore, comparison between the two communities of the inter-laboratory differences at T_2 versus T_1 suggests that this potential exaggeration of the annual decrement in FEV_1 was slightly greater in Long Beach (−13 ml/year, SE ± 7 ml/year) than in Lancaster (−2 ml/year, SE ± 7 ml/year). These decrements were not statistically different. Further, these differences observed in the 3 percent probability sample may not have existed in the entire group tested. The response rate for retesting of the 3 percent probability sample at T_2 was less than 50 percent.

Changes in Lung Function Tests

The means of the individual changes in tests are shown in Table 6A, for males and females 25–59 years of age, and stratified into five age groups for males (Table 6B) and females (Table 6C) 7–24 years of age at baseline. Non-adults were stratified to separate the growth phase of childhood from the declining trend in lung function which begins in late adolescence, as we have reported previously.[33] The levels of significance of differences between results in the two areas are shown in Table 7 for those age-and-sex-specific groups in which the probability was less than .05.

Changes for $\Delta N_{2_{750-1250}}$ and the spirometric indices except V_{75} were significantly worse in Long Beach participants 25–59 years. The level of significance for the difference in baseline and follow-up test results was highest for the single-breath nitrogen test. The levels of signifi-

TABLE 5—Means and Standard Errors of Spirometric and $\Delta N_{2750-1250}$ Tests at Baseline for Never Smokers Tested and Not Retested*

Test/Unit/ Residence	Retested			
	Lung Function Tests	Questionnaire only	Died	Not Retested
Number†				
Lancaster	829	311	19	740
Long Beach	499	156	18	516
FEV_1 (0/P%)**				
Lancaster	100 ± 0.5	100 ± 1.0	87 ± 4.8	99 ± .84
Long Beach	103 ± 0.7	101 ± 1.6	90 ± 6.5	101 ± 1.05
$FEF_{25-75\%}$(0/P%)**				
Lancaster	92 ± 1.7	96 ± 2.6	100 ± 6.8	91 ± 2.21
Long Beach	89 ± 2.0	94 ± 4.3	91 ± 11.7	85 ± 2.99
$\dot{V}_{50}$ (1/sec)				
Lancaster	4.1 ± 0.06	4.3 ± 0.09	4.4 ± 0.40	4.2 ± .08
Long Beach	4.1 ± 0.07	4.4 ± 0.12	4.4 ± 0.49	4.1 ± .12
$\dot{V}_{75}$ (1/sec)				
Lancaster	1.7 ± 0.03	1.7 ± 0.04	1.8 ± 0.18	1.7 ± .04
Long Beach	1.5 ± 0.04	1.7 ± 0.07	1.6 ± 0.22	1.4 ± .06
$\Delta N_{2750-1250}$ (%)				
Lancaster	0.71 ± 0.03	0.75 ± 0.04	1.33 ± 0.18	.79 ± .09
Long Beach	0.94 ± 0.04	0.73 ± 0.05	0.81 ± 0.17	.91 ± .05

*White, non-Spanish-surnamed only; 7–59 years of age at baseline; had not changed job or residence because of a respiratory problem.
†Minimum number in groups.
**Expected values derived from previously reported equations (reference 28).

TABLE 6A—Means of Annual Change in Lung Function Tests Results between Baseline and Retest 5–6 Years Later in Participants* 25–59 Years Who Never Smoked

	FEV_1	FVC	$FEF_{25-75\%}$	$\dot{V}_{50}$	$\dot{V}_{75}$	$\Delta N_{2750-1250}$
Males						
Lancaster (N = 132)	$-51^{\pm 5}$**	$-57^{\pm 7}$	$-54^{\pm 12}$	$-75^{\pm 21}$	$-44^{\pm 9}$	$-1.9^{\pm 1.2}$
Long Beach (N = 120)	$-66^{\pm 4}$	$-83^{\pm 6}$	$-93^{\pm 10}$	$-121^{\pm 14}$	$-53^{\pm 8}$	$3.2^{\pm 0.9}$
Females						
Lancaster (N = 259)	$-36^{\pm 3}$	$-41^{\pm 4}$	$-56^{\pm 7}$	$-83^{\pm 9}$	$-40^{\pm 7}$	$-1.5^{\pm 0.8}$
Long Beach (N = 188)	$-50^{\pm 3}$	$-61^{\pm 4}$	$-78^{\pm 7}$	$-111^{\pm 8}$	$-40^{\pm 5}$	$6.6^{\pm 1.2}$

*White, non-Spanish-surnamed only, had not changed job or residence because of a respiratory problem. FEV^1 exists at both times.
**Mean of annualized 5-year change; mean = (T2-T1/elasped months × 12 ± standard error).

cance of the differences were higher for most tests for adult women than for men except for the $FEF_{25-75\%}$.

Among non-adults, a significantly less favorable change in the results of the single-breath nitrogen test was observed in the youngest age group (7–10 years) for both males and females in Long Beach. A less favorable change in the results of the single-breath nitrogen test was observed in every age group under 25 years in Long Beach, but was not statistically significant in the 11–14 and 19–24 year groups among males and in the 15–18 year old group in females.

Among non-adult females the changes in spirometry were worse among Long Beach participants in every age group except for the $\dot{V}_{75}$ in the 7–10 year group and reached statistical significance for at least one spirometric test in every age stratum (Table 6C). Among males under 15 years, however, spirometry was not consistently worse in Long Beach participants. Above 15 years spirometry was consistently worse in Long Beach male participants, but did not reach statistical significance in the 15–18 year group. Above 19 years the results were significantly worse in Long Beach participants for all the spirometric indices except $\dot{V}_{75}$.

In every instance in which a statistically significant difference was observed the rate of change was worse among residents of the more polluted area, Long Beach. These included tests believed to reflect both small and large airways function. A significantly greater rate of deterioration was observed at an earlier age among females than among males.

Discussion

Although cohort studies provide the only absolute measures of risk in human populations, a number of problems arise in successfully completing them. Probably the most important is to achieve follow-up of a high proportion of participants. The proportion of all living participants completely retested in this study was low (47 percent and 45 percent in Lancaster and Long Beach, respectively). The major reason for not being retested, however, was not refusal, but migration out of the study area (39 percent and 47 percent, respectively.) The proportion completely retested of those remaining in the study areas was relatively high (75 percent and 77 percent). In the interval between baseline and retesting only a small percentage of participants moved

TABLE 6B—Means of Annual Change in Lung Function Test Results Between Baseline and Retest 5–6 Years Later in Male Participants[*] 7–24 Years Who Never Smoked

Males	FEV_1 (ml)	FVC (ml)	$FEF_{25-75\%}$ (ml/sec)	$\dot{V}_{50}$ (ml/sec)	$\dot{V}_{75}$ (ml/sec)	$\Delta N_{2_{750-1250}}$ (% × 100)
7–10 Years						
Lancaster (N = 96)	$287^{\pm 11}$[†]	$338^{\pm 14}$	$297^{\pm 18}$	$305^{\pm 19}$	$154^{\pm 14}$	$-6.7^{\pm 1.3}$
Long Beach (N = 32)	$286^{\pm 16}$	$327^{\pm 17}$	$307^{\pm 29}$	$320^{\pm 30}$	$181^{\pm 22}$	$-0.4^{\pm 1.5}$
11–14 Years						
Lancaster (N = 82)	$309^{\pm 11}$	$351^{\pm 14}$	$338^{\pm 19}$	$331^{\pm 21}$	$172^{\pm 18}$	$-1.5^{\pm 1.0}$
Long Beach (N = 14)	$313^{\pm 25}$	$352^{\pm 32}$	$335^{\pm 30}$	$329^{\pm 44}$	$182^{\pm 24}$	$-0.6^{\pm 2.1}$
15–18 Years						
Lancaster (N = 26)	$64^{\pm 17}$	$75^{\pm 17}$	$48^{\pm 30}$	$50^{\pm 36}$	$-15^{\pm 36}$	$-4.0^{\pm 2.4}$
Long Beach (N = 16)	$37^{\pm 14}$	$39^{\pm 18}$	$0^{\pm 23}$	$-10^{\pm 27}$	$-20^{\pm 14}$	$3.2^{\pm 2.1}$
19–24 Years						
Lancaster (N = 21)	$-16^{\pm 13}$	$-12^{\pm 14}$	$-8^{\pm 20}$	$0^{\pm 32}$	$-45^{\pm 21}$	$-1.8^{\pm 2.4}$
Long Beach (N = 23)	$-70^{\pm 13}$	$-72^{\pm 15}$	$-125^{\pm 26}$	$-149^{\pm 40}$	$-90^{\pm 18}$	$0.4^{\pm 0.9}$

[*] White, non-Spanish-surnamed only, had not changed job or residence because of a respiratory problem. FEV_1 exists at both times.

[†] Mean of annualized 5-year change; mean = (T2-T1/elapsed months × 12 standard error).

TABLE 6C—Means of Annual Change in Lung Function Test Results between Baseline and Retest 5–6 Years Later in Female Participants[*] 7–24 Years Who Never Smoked

Females	FEV_1 (ml)	FVC (ml)	$FEF_{25-75\%}$ (ml/sec)	$\dot{V}_{50}$ (ml/sec)	$\dot{V}_{75}$ (ml/sec)	$\Delta N_{2_{750-1250}}$ (% × 100)
7–10 Years						
Lancaster (N = 68)	$236^{\pm 7}$[†]	$261^{\pm 9}$	$269^{\pm 16}$	$258^{\pm 19}$	$125^{\pm 13}$	$-7.9^{\pm 1.7}$
Long Beach (N = 29)	$209^{\pm 10}$	$216^{\pm 15}$	$258^{\pm 27}$	$232^{\pm 25}$	$146^{\pm 22}$	$+3.7^{\pm 1.8}$
11–14 Years						
Lancaster (N = 85)	$124^{\pm 9}$	$120^{\pm 11}$	$147^{\pm 17}$	$118^{\pm 19}$	$98^{\pm 13}$	$-1.7^{\pm 1.2}$
Long Beach (N = 24)	$68^{\pm 16}$	$75^{\pm 22}$	$62^{\pm 25}$	$41^{\pm 23}$	$6^{\pm 25}$	$+5.5^{\pm 2.0}$
15–18 Years						
Lancaster (N = 49)	$11^{\pm 7}$	$23^{\pm 10}$	$-26^{\pm 18}$	$-28^{\pm 20}$	$-48^{\pm 18}$	$-0.4^{\pm 1.5}$
Long Beach (N = 16)	$-20^{\pm 9}$	$-16^{\pm 10}$	$-80^{\pm 21}$	$-82^{\pm 21}$	$-66^{\pm 17}$	$+4.4^{\pm 2.0}$
19–24 Years						
Lancaster (N = 27)	$-6^{\pm 10}$	$-11^{\pm 13}$	$-30^{\pm 20}$	$-20^{\pm 25}$	$-53^{\pm 17}$	$-2.9^{\pm 2.1}$
Long Beach (N = 36)	$-18^{\pm 13}$	$-18^{\pm 17}$	$-60^{\pm 15}$	$-77^{\pm 16}$	$-64^{\pm 14}$	$+2.9^{\pm 1.2}$

[*] White, non-Spanish-surnamed only, had not changed job or residence because of a respiratory problem. FEV_1 exists at both times.

[†] Mean of annualized 5-year change; mean = (T2-T1/elapsed months × 12 ± standard error).

because of respiratory problems (2.1 percent in Long Beach and 0.5 percent in Lancaster).

A major way of assessing the impact of a lower-than-desired response rate is to determine the differences in the baseline test results between participants who were retested and those who were not. The difference in the mean test indices at baseline were not significantly different in those who were not retested than in those who subsequently completed lung function retesting, suggesting that those who were not retested were substantially similar in respiratory function to those who were.

A particular problem for cohort studies of the relationship of air pollution to changes in lung function test results is the possible impact of air pollution levels on the day of testing.[8] For that reason, analysis was carried out comparing the results of lung function testing with levels of each of the major pollutants on the day of testing. No significant correlations with test results were seen and no consistent pattern emerged for the major pollutants occurring at higher concentrations in Long Beach, suggesting that the range of levels of these pollutants prevailing during the testing periods had relatively little impact on the results of the lung function testing in this study.

The tests used in this study can be affected by a number of confounding factors including instrument characteristics, subject effort, and technician performance.[29] Thus, there is continual concern with the accuracy and reliability of the equipment and testing procedures, particularly in a mobile laboratory in which movement of the equipment from site to site can precipitate instrument drift. For that reason, the test-retest difference in FEV_1 observed in the five to six year interval in the Mobile Lung Function Laboratory was compared with results obtained on a 3 percent subsample of participants tested on both occasions at the UCLA Reference Laboratory. This comparison revealed only a small difference between the two laboratories in favor of showing a slightly greater deficit in FEV_1 in Long Beach if, in fact, the difference in deficit was also present in the entire group tested.

For the lung function tests, the mean of the absolute change in individual test results was used rather than differences in predicted values. Thus, differences between communities in age and height could have biased the results. The mean height was, however, similar in the two

TABLE 7—Magnitude of t* and Significance Levels for Differences in Rate of Change of Lung Function Test Results in Lancaster and Long Beach Participants with $p < .05$

Sex/Age (years)		FEV_1	FVC	$FEF_{25-75\%}$	$\dot{V}_{50}$	$\dot{V}_{75}$	$\Delta N_{2750-1250}$
Males							
7–10	T=	—	—	—	—	—	−3.12
	P<	—	—	—	—	—	.002
11–14	T=	—	—	—	—	—	—
	P<	—	—	—	—	—	—
15–18	T=	—	—	—	—	—	−2.25
	P<	—	—	—	—	—	.03
19–24	T=	2.90	2.93	3.89	3.11	—	—
	P<	.006	.005	3×10^{-4}	.003	—	—
25–59	T=	2.15	2.74	2.55	1.97	—	−3.64
	P<	.032	.006	.011	.049	—	3×10^{-4}
Females							
7–10	T=	2.16	2.63	—	—	—	−4.85
	P<	.033	.001	—	—	—	1×10^{-8}
11–14	T=	2.87	—	2.52	2.04	3.33	−3.13
	P<	.005	—	.013	.044	.002	.002
15–18	T=	2.24	—	2.18	—	—	—
	P<	.023	—	0.33	—	—	—
19–24	T=	—	—	—	1.99	—	−2.66
	P<	—	—	—	.051	—	.01
25–59	T=	3.48	3.49	2.20	2.37	—	−6.20
	P<	5×10^{-4}	5×10^{-4}	.028	.018	—	1.6×10^{-9}

$$^{*}t = \frac{\bar{X}_1 - \bar{X}_2}{S} \text{ where } S = \frac{S_1^2(N_1-1)+S_2^2(N_2-1)}{N_1+N_2-2}\left(\frac{1}{N_1}+\frac{1}{N_2}\right)$$

areas. The mean ages of the two populations were similar although the females in Long Beach were on average three years older than females in Lancaster, whereas the males in Long Beach were on average one year younger than in Lancaster. A prior study of this population suggested that the rate of decline in FEV_1 remained fairly constant over the age range 25–59 years.[33]

Recent reports have suggested that young individuals living in homes supplied with gas, especially for cooking, may have more respiratory problems than persons in homes that are supplied with electricity.[34–38] The percentage of homes that used gas for heating, however, was exactly the same in both communities (94 percent).

Many people in Southern California commute long distances to workplaces in which they may experience air pollution patterns different from those at their place of residence. Individuals in Long Beach commuted to areas of lower levels of the major pollutants (except for oxidants), whereas most residents of Lancaster who commuted did so to areas of higher levels of pollution.[28] This pattern would tend to minimize the actual difference in pollutant exposures for participants in the two study areas and, thus, would reduce the likelihood of observing a pollution effect on lung function test results.

A number of other factors that could have biased the results have been dealt with by restricting the population analyzed in this paper. Differences in racial distribution were reduced by including only White non-Spanish-surnamed participants in the analyses and by excluding individuals who reported a history of changing residence or occupation because of a respiratory problem at baseline. The number of individuals who had worked in an occupation that might be associated with respiratory injury was too high to permit excluding them from analysis. Their occupations were not concentrated in any one or two types of exposure and the duration of their employment in any type was generally less than five years. The proportion with such a history, however, was higher in Lancaster than in Long Beach and, thus, including them in the analysis probably reduced the likelihood of observing less favorable changes in Long Beach.

The average annual change in FEV_1 among these never-smoking adults was higher than has been reported in most previous population studies.[6–8,39–41] The levels observed in our study are comparable with some of the more recent reports of changes in lung function over time in occupational groups.[42–44] Although we looked intensively for methodologic problems that may have caused these large rates of decline in adults, we were unable to identify any consistent factors that would account for the magnitude of decline observed in FEV_1. One possibility is that these large average annual declines in FEV_1 may reflect the accumulated respiratory effects of chronic exposure to air pollution. Supporting this hypothesis is the previous observation that the annual rate of decline in FEV_1 between 25 and 59 years of age in these populations increased only very slightly with age[33] in contrast to the significant increase in the annual rate of decline with aging reported in other publications.[45,46] Because of local migration patterns, fewer of our older participants and more of our younger participants were raised in the Los Angeles area. This, and the consistently poorer (more positive) changes in $\Delta N_{2750-1250}$, even in the youngest participants in Long Beach, suggest that the younger members of our cohorts raised in Los Angeles may experience greater drops in their lung function test parameters when they reach later life than the drop currently observed in the older members of our cohorts who were raised in less polluted areas. If in fact this hypothesis is true, it is cause for concern.

The test demonstrating differences most frequently and with the highest levels of significance was the $\Delta N_{2_{750-1250}}$. Furthermore, in the youngest age groups this test often demonstrated the only significantly worse deterioration, suggesting that physiologic changes may be occurring in children that are not as readily identified by the presumably less sensitive spirometric tests until adulthood. Since the $\Delta N_{2_{750-1250}}$ is thought to reflect function primarily in the small airways, the results of this cohort study suggest that this may be the site of earliest injury due to exposure to the type of air pollutants found in the Long Beach area. Buist and colleagues[47] and McCarthy, *et al*,[48] have demonstrated that the single-breath nitrogen test is more sensitive than spirometry in detecting early functional abnormality involving the small airways. In cross-sectional studies we were not able to observe differences between communities in this parameter.[27,28,30] We had previously observed, however, that the $\Delta N_{2_{750-1250}}$ is subject to considerable variability which may account for the difficulty in identifying significant differences in cross-sectional studies.[49] Inter-subject variability may be a greater problem in cross-sectional studies than intra-subject variability in cohort studies in which differences in two measurements from the same individual are being considered.

Among adults who never smoked, rate of decline in all of the tests noted in Table 7 was significantly worse in residents of Long Beach except for the $\dot{V}_{75}$ in which a directionally similar difference did not attain statistical significance.

Although the large differences in the results of $\Delta N_{2_{750-1250}}$ in children in even the youngest age groups suggest that the small airways are the earliest site of pathology resulting from chronic exposure to the mix of air pollutants occurring in Long Beach, the greater deterioration in the spirometric indices indicates that by the time residents of the area reach adulthood, the adverse changes in the small airways have progressed and/or the large airways have become affected as well. It remains to be determined if the adverse changes in the lungs resulting from chronic exposure to this mix of pollutants causes sufficient respiratory impairment over a lifetime to reduce their ability to function satisfactorily or only reduces their considerable reserve capacity without causing clinically significant functional impairment.

Ware, *et al*, have studied the respiratory symptom and lung function effects of chronic exposure to TSP, SO_2, and TSO_4 at two different examinations one year apart in children 6–10 years of age in six cities in the United States and have found no relationship between respiratory parameters and levels of these pollutants.[50] While it is tempting to compare the results of that study to the present study there are several problems with doing so. First, the measures of air pollutants were different. Whereas Ware, *et al*, used the mean of the 24-hour values, we used the means of the daily peak hourly values for SO_2, although for TSP and TSO_4 we used the 24-hour totals. Of greater importance is the fact that we looked at the mean annual change in lung function values within each individual over a five to six year interval whereas Ware, *et al*, looked at the correlations of the pollutant values with FEV_1 and FVC at two time periods only one year apart. Thus, their study design was essentially two consecutive cross-sectional correlations to mean lung function values for groups of children in each area whereas we looked at the mean of change within each individual over a five-year period. Nonetheless, there are some similarities between the studies. Neither found consistent correlations between respiratory symptoms and levels of air pollutants. Further, we did not observe a significant correlation between change in lung function and pollutant level in males 7–10 years of age. The correlation we observed in both males and females 7–10 years was in the single-breath nitrogen test, a possibly more sensitive measure of early decrement in lung function. It will be interesting to observe whether Ware, *et al*, find correlations between pollutant levels and lung function with additional follow-up of the children in these six cities. We did not consistently observe significant differences in FEV_1 and FVC in boys until 15–18 years. We did, however, observe consistent difference in earlier age groups using the single-breath nitrogen test.

We previously observed a greater deficit in several lung function test parameters between Lancaster and an area subject to very high levels of oxidants, Glendora.[26,30] The greatest difference in pollutant exposures between Long Beach and Glendora was in the levels of oxidants (much higher in Glendora), nitrogen dioxide (higher in Long Beach), and SO_2 (higher in Long Beach). Mean levels of SO_4 were similar in these two communities and higher than in Lancaster which had primarily oxidant exposure, which, although high, was half that in Glendora. In the Long Beach comparison, the measures incorporating the large airways (FEV_1, FVC) were consistently significantly different in adults whereas only the FEV_1 in adult females in the Glendora comparison were significant. On the other hand, a significant difference in $\dot{V}_{75}$ was observed only for Glendora. The deterioration in the single-breath nitrogen test in Glendora was significantly worse in every age-sex group under 25 years compared whereas they were significantly worse only in males 7–10 years and 15–18 years and in females 7–14 and 19–24 years in Long Beach. Significant differences in most of the spirometric indices in these younger age groups were not consistently identified in either polluted community compared to Lancaster. It is possible that the type of pollutants that were present in greatest concentrations in Long Beach (NO_2, hydrocarbons, and SO_2) were more likely to cause damage in the large airways whereas oxidants were more likely to cause damage in the smaller airways as reflected by the greater magnitude of differences for the $\Delta N_{2_{750-1250}}$ and $\dot{V}_{75}$ in Glendora. Further, exposure to oxidant pollutants may cause measurable damage earlier (according to the single-breath nitrogen test) than SO_2 and NO_2. These interpretations are consistent with results of controlled exposure studies in experimental animals which demonstrate that characteristic ozone-induced lesions occur in the lung periphery (centiacinar region)[51,52] while chronic exposure to SO_2 produces clinical, physiologic, and histologic changes of chronic bronchitis with greater involvement of proximal than distal airways.[53]

Given the nature of epidemiological studies which must deal with populations as they actually exist in the community rather than with laboratory animals which can be kept under experimental conditions, these inferences should be investigated further. It appears, however, that chronic exposure to either mix of pollutants results in less rapid growth of lung function in children and a greater rate of deterioration in adulthood.

These observations have serious implications for safeguarding the health of the public and controlling levels of pollutants in urban areas in the United States. These studies need to be verified by others, but the implications for health should not be ignored by legislators and government administrators. Given the difficulties of carrying out epidemiologic studies of cohorts of individuals over long periods of time the probability of being able to observe statistically significant differences is smaller than for similarly designed studies done under laboratory conditions. For this reason, the burden of proof should now be to prove that high levels of these pollutants do not affect lung function and, therefore, until that is demonstrated efforts should be increased to improve the quality of air in the urban areas of the United States. ☐

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APPENDIX A

Single-breath Nitrogen Test

Ventilation distribution was determined using an Ohio-Series 700 nitrogen analyzer, Electro-Medical (Ohio Medical Products, Houston) model 780 spirometer, electronics module, samplehead assembly, remote vacuum pump assembly, and Hewlett-Packard X-Y recorder (Hewlett-Packard, San Diego) with time base.

Subjects performed these tests in the standing position with a nose clamp and flanged rubber mouthpiece in place. After a few normal tidal breaths and stabilization of the system, the subject was asked to exhale completely and to hold his breath briefly at the residual volume position. The valve was switched to the reservoir containing 100 percent oxygen and the expiratory port simultaneously opened to the (electronic) spirometer. The subject was asked to inhale fully and deeply, within less than 4 seconds, to total lung capacity, then to exhale slowly and completely at a constant flow rate maintained between 0.5 and 0.8l/s according to an X-Y tracing which was visible to him. Residual volume was reached over no longer than 10 s. The expired volume and expired nitrogen concentrations were simultaneously recorded on the horizontal and vertical axes of the X-Y recorder. A minimum of three trials was completed for each resident.

Each tracing was read by only one individual, the Chief Technologist, who assigned values for the percentage of nitrogen (percent N_2) in the expired air at 750 and 1,250 ml of expirate and identified both the point at which the inflection between phase III and Phase IV occurred and the point at which total volume was exhaled (residual volume).

A Measles Outbreak at a College with a Prematriculation Immunization Requirement

Bradley S. Hersh, MD, MPH, Lauri E. Markowitz, MD, Richard E. Hoffman, MD, MPH, Daniel R. Hoff, PA, Mary J. Doran, Jessica C. Fleishman, Stephen R. Preblud, MD, and Walter A. Orenstein, MD

ABSTRACT

Background. In early 1988 an outbreak of 84 measles cases occurred at a college in Colorado in which over 98 percent of students had documentation of adequate measles immunity (physician diagnosed measles, receipt of live measles vaccine on or after the first birthday, or serologic evidence of immunity) due to an immunization requirement in effect since 1986.

Methods. To examine potential risk factors for measles vaccine failure, we conducted a retrospective cohort study among students living in campus dormitories using student health service vaccination records.

Results. Overall, 70 (83 percent) cases had been vaccinated at ≥12 months of age. Students living in campus dormitories were at increased risk for measles compared to students living off-campus (RR = 3.0, 95% CI = 2.0, 4.7). Students vaccinated at 12–14 months of age were at increased risk compared to those vaccinated at ≥15 months (RR = 3.1, 95% CI = 1.7, 5.7). Time since vaccination was not a risk factor for vaccine failure. Measles vaccine effectiveness was calculated to be 94% (95% CI = 86, 98) for vaccination at ≥15 months.

Conclusions. As in secondary schools, measles outbreaks can occur among highly vaccinated college populations. Implementation of recent recommendations to require two doses of measles vaccine for college entrants should help reduce measles outbreaks in college populations. (*Am J Public Health* 1991;81:360–364)

Introduction

Since licensure of live measles vaccine in the United States in 1963, there has been a marked reduction of reported measles incidence. In 1978, the United States began an effort to eliminate indigenous measles. The elimination strategy called for achieving and maintaining high immunization levels, careful surveillance, and aggressive outbreak control. The high immunization levels were to be achieved using a single dose of measles vaccine.[1]

Although there has been a 98 to 99 percent reduction in the number of annual reported cases from the pre-vaccine era, measles outbreaks still occur. Two predominant patterns of measles transmission have been observed: outbreaks occurring in predominately unvaccinated inner-city preschool-age children, and outbreaks occurring in highly vaccinated secondary school and college populations.[2,3] While most outbreaks occur in secondary schools, cases from college campuses have accounted for between 1.5 percent and 19.8 percent of cases each year from 1980 to 1989 (Table 1).[4,5] In 1983, the record low year for reported measles cases, campus outbreaks and non-campus cases associated with these outbreaks accounted for 32.4 percent of total cases.[6]

There are several reasons why measles outbreaks may be likely to occur in colleges and universities. Some college-aged students may not have been vaccinated in the early years following licensure of measles vaccine; others may not have been appropriately vaccinated at school entry or thereafter. Data from serological surveys suggest that between 5 and 15 percent of college-aged individuals may be susceptible to measles.[7] The propensity for college students to congregate in large groups and their travel to areas where measles transmission is endemic may increase the potential for introduction and spread of measles virus on college campuses.

School immunization laws have been demonstrated to be an effective means of assuring high measles vaccination levels and preventing measles outbreaks among kindergarten through 12th grade children and adolescents.[8] The American College Health Association[9–11] and the Immunization Practices Advisory Committee of the US Public Health Service (ACIP)[12] have recommended that colleges and universities require students to present evidence of measles immunity as a condition for enrollment. In early 1988, a measles outbreak occurred at a college that had implemented such an immunization requirement. This outbreak provided an opportunity to investigate the occurrence of measles in a highly vaccinated college population and to examine potential risk factors for measles vaccine failure.

Address reprint requests to Bradley S. Hersh, MD, MPH, Division of HIV/AIDS, Center for Infectious Diseases, Centers for Disease Control, Atlanta, GA 30333. Ms. Fleishman is with the University of Colorado Health Sciences Center; Drs. Markowitz and Orenstein are also with the Division of Immunization at CDC; Dr. Preblud, now deceased, was also with that Division; Dr. Hoffman and Ms. Doran are with the Colorado Department of Health, Denver; Mr. Hoff is a Physician Assistant with the Student Health Service, Fort Lewis College, Durango, CO. This paper, submitted to the Journal February 26, 1990, was revised and accepted for publication August 13, 1990.

TABLE 1—Reported Measles Cases on US College Campuses, 1980–89

Year	Total US Cases	Campus Cases No.	Campus Cases %	No. of Colleges Reporting Cases
1980	13,506	200	(1.5)	36
1981	3,124	101	(3.2)	19
1982	1,714	115	(6.7)	14
1983	1,497	296	(19.8)	19
1984	2,587	67	(2.6)	17
1985	2,822	354	(12.6)	26
1986	6,282	174	(2.8)	21
1987	3,655	141	(3.9)	22
1988	3,396	267	(7.9)	20
1989	18,193	1,643	(9.0)	93

Background

Fort Lewis College is a public four-year college in Durango, Colorado. There were 3,555 students registered at the time of the outbreak. Since September 1986, the college has required students to provide proof of immunity to measles, as well as to other vaccine-preventable diseases, before enrolling. Adequate measles immunity was considered documentation of physician-diagnosed measles, receipt of live measles vaccine on or after the first birthday, or serologic evidence of immunity. Prior to this outbreak almost 99 percent of the student population were in compliance with this requirement and had documentation on file at the student health center.

Methods

Case Definition and Case Ascertainment

A case of measles was defined as an illness that met the Centers for Disease Control clinical case definition[13] (generalized maculopapular rash ≥3 days duration, fever ≥101 F, if measured, and at least one of the following: cough, coryza, or conjunctivitis). A suspected case was any rash illness with fever; a confirmed case met the clinical case definition and was either serologically confirmed or epidemiologically linked to another confirmed or clinical case of measles. A case was considered to be serologically confirmed if there was a fourfold rise in titer between acute and convalescent specimens by the complement-fixation assay.

Students were informed about the outbreak by class announcements, articles in the student newspaper, and reports on the campus radio station. They were instructed to report to the student health service at the onset of any febrile or rash illness. A disease control specialist from the Colorado Department of Health investigated every suspected case.

Outbreak Control

On January 8, 1988, after the third clinical case of measles was reported at the college, a manual review of college immunization records was conducted. A total of 40 students were found without documentation of measles immunity. All but three of these students were promptly vaccinated. The students not vaccinated had medical exemptions to vaccination and were excluded from campus for the duration of the outbreak. On January 12, 1988, the Colorado Department of Health placed the campus under public health orders. All indoor sporting events and other indoor college activities on or away from campus were restricted to participants who could show proof of measles immunity. All Fort Lewis participants in intercollegiate activities were examined prior to an event for signs and symptoms of measles prodrome. Furthermore, no spectators were allowed to attend these events. In order to prevent further measles transmission, all suspected measles cases among Fort Lewis College students were isolated in special dormitory rooms located near the student health service.

Retrospective Cohort Study

To investigate risk factors for measles vaccine failure, we conducted a retrospective cohort study among all students living in campus dormitories. Vaccination histories were obtained from student health center vaccination records. Student variables collected include: name, sex, race, class, date of birth, and housing status. Students who had received measles vaccine less than 14 days prior to rash onset were considered to be unvaccinated. We investigated the following potential risk factors for vaccine failure: number of doses of measles vaccine received, age at vaccination, and time since vaccination. Since several investigations of recent measles outbreaks have found individuals vaccinated before 1980 to be at higher risk for measles than those vaccinated in 1980 or after,[14] we also evaluated year of vaccination.

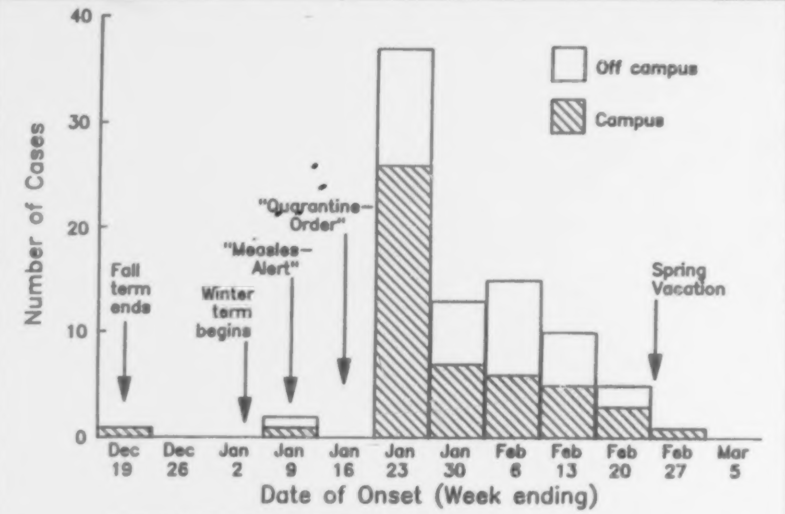
Age at vaccination was defined as the number of months between the date of birth and the most recent vaccination date on the vaccination record. Time since vaccination was calculated as the number of years between the date of most recent vaccination and the beginning of the outbreak (January 1, 1988). The age-at-vaccination analysis was limited to single dose vaccinees who were vaccinated before 1980; the time-since-vaccination and year-of-vaccination analyses were limited to single dose vaccinees who were vaccinated at 15 months of age or older. Finally, in order to verify information contained in case vaccination records, efforts were made to contact health care providers who had administered the vaccine. A physician-verified vaccination record was defined as confirmation of at least month and year of vaccination.

Differences in attack rates among groups in the retrospective cohort study were measured using the chi-square and Fisher's exact tests. Measles vaccine effectiveness and 95 percent confidence intervals were calculated using the method described by Orenstein, et al.[15]

Results

Between December 15, 1987, and March 10, 1988, a total of 84 measles cases in five generations of measles transmission occurred among Fort Lewis College students (Figure 1). All cases satisfied the clinical case definition for measles. Of the total cases, 11 (13 percent) were serologically confirmed.

The index case was an 18-year-old female freshman from South Carolina who had been vaccinated against measles at 14 months of age. The source of her infection is unknown. Measles transmission from the index case to second generation campus cases occurred at the Durango airport. Further measles transmission occurred primarily in campus dining halls, in classrooms, and at several campus social events. There was no spread of measles to other schools despite 10 inter-collegiate indoor activities or games, and only five cases were reported among non-college



Source: Colorado Department of Health

FIGURE 1—Reported Measles Cases by Week of Rash Onset, Fort Lewis College, December 1987–March 1988.

residents of Durango. Two of the non-college cases were directly linked to the index case.

Of the 84 cases, 70 (83 percent) occurred in individuals with evidence of measles vaccination at 12 months of age or older, including 20 (24 percent) vaccinated at 12 to 14 months of age. Of the 70 cases, vaccination histories (89 percent) were verified by the physician who had administered the vaccine. Five cases (6 percent) occurred in persons vaccinated before the first birthday, and 9 cases (11 percent) occurred in persons without documentation of measles vaccination. Of the cases in students without documentation of measles vaccination, vaccine was not routinely recommended for four (two born before 1957, two with a history of physician-diagnosed measles), and a fifth student had a religious exemption to vaccination.

Forty-six cases (55 percent) were freshmen, 46 (55 percent) were males, and 53 (63 percent) occurred in students living in campus dormitories. Seventy-five (89 percent) cases were White, seven (8 percent) were Native American, one case (1 percent) was Black, and one (1 percent) was Hispanic. Cases ranged in age from 17–32 years; 62 (74 percent) occurred in persons 18–20 years of age.

The overall attack rate among Fort Lewis College students was 2.4 percent. There were no differences in attack rates by race. Attack rates by housing status, sex, and class are summarized in Table 2. Students living in campus dormitories were at higher risk for measles than students living off-campus. When the analysis was limited to students living in campus dormitories, differences in attack rates by class are no longer observed.

In the retrospective cohort study among students living in campus dormitories, students without documentation of measles vaccination and those vaccinated before 12 months of age were at greater risk for measles than students vaccinated at 12 months of age or older (Table 3). No case of measles occurred in 46 students who had received two doses of measles vaccine after 12 months of age. Students vaccinated once at less than 12 months or at 12 to 14 months of age were at greater risk for measles than those vaccinated once at 15 months of age or greater. Although students vaccinated prior to 15 months of age were at increased risk for measles, these cases accounted for only 30 percent of the total cases. Students vaccinated before 1980 were not at significantly increased risk for measles compared with those vaccinated in 1980 or after. No trend for increasing attack rate with increasing time since vaccination was observed among students vaccinated once at 15 months of age or older (Figure 2).

The estimates of vaccine effectiveness among students living in campus dormitories was 80 percent (95 percent CI = 51–92 percent) for vaccination at 12 to 14 months of age and 94 percent (95 percent CI = 86–98 percent) for vaccination at 15 months or greater.

Discussion

The Fort Lewis College measles outbreak occurred in spite of virtually complete implementation of a prematriculation measles vaccination requirement. It was thus similar to outbreaks that have occurred in recent years among highly vaccinated junior and senior high school populations.[16–21] While the prematriculation requirement may have prevented a larger outbreak, it is apparent that measles outbreaks, such as the one described, can occur in highly vaccinated populations.

In this outbreak, living in campus dormitories was found to be a risk factor for acquiring measles. The relatively crowded living conditions and high contact rate among students in a college dormitory can facilitate measles virus transmission from infected to susceptible individuals.

Consistent with other investigations,[22] vaccination before the age of 15 months was found to be a risk factor for measles vaccine failure. The increased relative risk observed in persons vaccinated before 15 months of age has been attributed to the persistence of maternal antibody in infants, which interferes with live measles vaccine. Although age of vaccination was found to be a risk factor, 60 percent of cases occurred in students vaccinated at 15 months or older. As in other investigations,[17,19,21,23,24] there was no evidence in this outbreak that the risk for measles increases with time since vaccination.

We did not find evidence that vaccination prior to 1980 was a risk factor for measles in this outbreak. Other investigations, however, have found this to be a risk factor for measles vaccine failure.[14] Prior to the addition of an improved heat stabilizer to measles vaccine in 1979, individuals may have received a less effective vaccine, causing lower rates of seroconversion and an increased risk for measles. Alternatively, one may speculate that the finding of a lower risk for measles in individuals vaccinated since 1980 may be due to the possibility that some of these persons may have, in fact, been receiving a second dose of measles vaccine. Students may often find it easier to receive a second dose of measles vaccine than to obtain documentation of their initial measles vaccination.

The investigation of this outbreak found measles vaccine to be over 90 percent effective for vaccination at 15 months or greater. Other investigations have found similar results.[25] The occurrence of

TABLE 2—Risk Factors for Measles, Attack Rates by Student Characteristics, Fort Lewis College

Characteristic	Students	Cases	Attack Rate	Relative Risk	95% CI
Housing					
Dormitories	1,278	53	4.1%	3.0	(2.0–4.7)
Off-campus	2,277	31	1.4%	1.0	referent
Sex					
Male	1,922	46	2.4%	1.0	(0.7–1.6)
Female	1,633	38	2.3%	1.0	referent
Class					
Freshman	1,347	46	3.4%	4.3	(1.9–10.1)
Sophomore	842	19	2.3%	2.9	(1.2–7.2)
Junior	603	13	2.1%	2.7	(1.1–7.2)
Senior	763	6	0.8%	1.0	referent
Class*					
Freshman	759	34	4.5%	2.5	(0.3–17.7)
Sophomore	352	13	3.7%	2.0	(0.3–15.2)
Junior	112	5	4.5%	2.5	(0.3–20.5)
Senior	55	1	1.8%	1.0	referent

*Among students living in campus dormitories.

measles outbreaks in populations with high immunization levels suggests that the estimated 5 to 10 percent vaccine failure rate may still provide enough susceptibles to allow sustained transmission of a virus as communicable as measles in some settings, such as a college dormitory.

The two national advisory bodies for immunization policy, the Committee on Infectious Diseases of the American Academy of Pediatrics (Red Book Committee) and the ACIP, have recently considered the problem of measles outbreaks in highly vaccinated populations and have recommended the institution of a routine two-dose measles vaccination schedule.[26,27] Furthermore, both committees recommend that college students have, at the time of college entry, documentation of having received two doses of measles vaccine after the first birthday. The American College Health Association has endorsed this recommendation.[28] Prematriculation immunization requirements provide a mechanism for colleges and universities to ensure that students are in compliance with this new recommendation.

Whether a second dose of measles vaccine will significantly increase population measles immunity and decrease the risk of measles outbreaks is not conclusively known. However, this investigation and several others have found lower risks for measles in persons who received two doses of vaccine.[17,19,21,23] The power of these studies to detect statistically significant differences, however, was low due to the relatively small number of multiple dose vaccinees. Experience from the military suggests that a second dose of measles vaccine should be effective in preventing outbreaks.[29]

Measles elimination has proved more difficult than previously predicted.[30] A single dose strategy, while highly effective, has not been able to totally prevent measles outbreaks. The full implementation of a routine two-dose measles vaccination schedule, however, should help reduce measles outbreaks in highly vaccinated populations. ☐

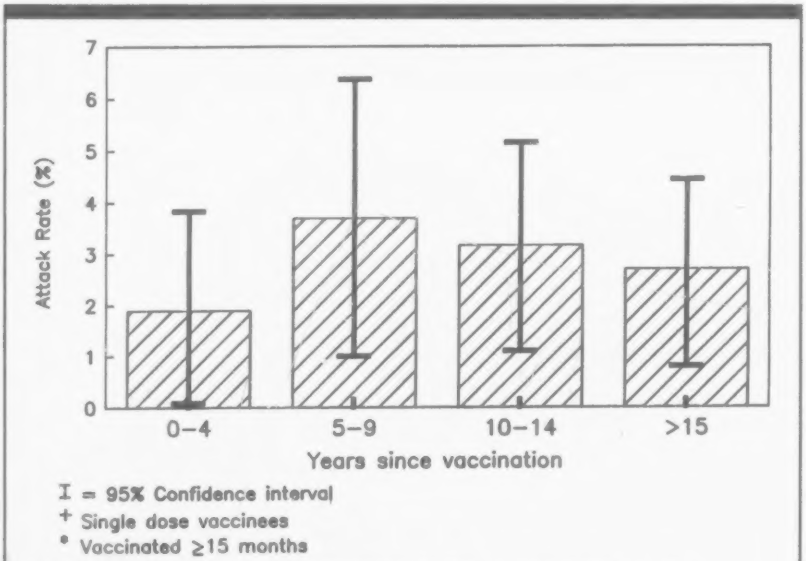


FIGURE 2—Attack Rate by Time Since Vaccination, Fort Lewis College.†*

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TABLE 3—Risk Factors for Measles for Students Living in Campus Dormitories, Fort Lewis College

Characteristic	Students	Cases	Attack Rate	Relative Risk	95% CI
Doses					
0 doses	6	3	50.0%	13.1	(5.6–30.8)
1 dose (<12 mos)	52	5	9.6%	2.5	(1.1–6.1)
1 dose (≥12 mos)	1,156	44	3.8%	1.0	referent
2 doses (1 dose <12 mos and 1 dose ≥12 mos)	18	1	5.6%	1.5	(0.2–10.0)
2 doses (both doses ≥12 mos)	46	0	0.0%	0.0	(0.0–1.7)
Age at vaccination[a,b]					
<12 mos	52	5	9.6%	3.0	(1.2–7.7)
12–14 mos	164	16	9.8%	3.1	(1.7–5.7)
≥15 mos	695	22	3.2%	1.0	referent
Year of vaccination[a,c]					
<1980	695	22	3.2%	1.6	(0.6–3.8)
≥1980	297	6	2.0%	1.0	referent

[a] Single dose vaccinees
[b] Vaccinated before 1980
[c] Vaccinated ≥ 15 months

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Assessment of Drug Abuser Treatment Needs in Rhode Island

William E. McAuliffe, PhD, Paul Breer, PhD, Nancy White Ahmadifar, PhD, and Cathie Spino, DSc

ABSTRACT

Background. Rhode Island's Division of Substance Abuse asked us to assess the State's drug treatment needs and make recommendations regarding its treatment system for the next three years.

Methods. We used a statewide telephone drug use survey of 5,176 households supplemented by drug-related hospital discharges, Division of Drug Control statistics, and interviews with providers, state officials, and out-of-state experts. Drug abuse was measured with items from the Diagnostic Interview Schedule. Abusers were asked if they were receiving or wanted to receive treatment.

Results. Survey responses, used to estimate the unmet need for drug treatment, indicated a need to triple drug treatment services. Regression models using survey data indicated that the treatment network was overly centralized in the Providence area. Interviews with state officials, clinicians, and out-of-state experts provided material for recommendations on reimbursement policy, treatment mix, quality assurance, and cost containment.

Conclusions. The RI Department of Health's certificate-of-need program adopted our overall recommendation for tripling the drug treatment system as its guideline in evaluating proposals for new treatment facilities. With State funding of a new adolescent center and expansion of outpatient slots in the private sector, this recommendation has now been fully implemented. (*Am J Public Health* 1991;81:365–371)

Introduction

Recent federal and foundation initiatives on drugs and acquired immunodeficiency syndrome (AIDS) among intravenous (IV) users have led many states and communities to assess their needs for drug treatment services. Up to now, attempts to determine drug abuse treatment needs have lacked adequate tools. In the 1970s when substantial public funds for drug treatment first became available, planners based their needs assessments on census data, arrests, treatment admissions, and occupancy rates.[1,2] At best, however, such statistics represent only indirect approximations of addict prevalence and may be influenced by actions of the agencies that report the data.[3] For example, if services are too expensive, not well located, or not well matched to addict clinical or cultural needs, underutilization of treatment facilities may occur even when there is substantial unmet need for services in a state. The present study sought to develop improved methodologies.

In 1986, the Rhode Island Division of Substance Abuse (DSA) sought documentation of unmet need for drug abuse services to convince legislators of the necessity for third-party reimbursement of drug abuse treatment and to establish guidelines for regulators evaluating proposed treatment facilities. DSA had previously contracted with Brown University for a needs assessment of the combined alcohol and drug treatment system.[4] Although the Brown Report estimated drug use prevalence in Rhode Island from age- and gender-specific *national* rates reported in a 1982 survey of drug use by the National Institute on Drug Abuse (NIDA), the Report's authors recommended that a local drug abuse prevalence survey be conducted.

Several developments in the last two decades have made improved needs assessments possible for state and local agencies. In 1972, the federal government demonstrated the feasibility of household surveys of nonmedical drug use employing face-to-face interviews with a cross-section of the nation's population.[5] Some states (e.g., New York and New Jersey) have now initiated periodic household surveys of their own.

In 1981 and 1986, the New York State Division of Substance Abuse Services conducted a statewide survey of drug use by telephone interview.[6] A pilot study (1980) found that random digit dialing and a sample drawn from local telephone directories both returned response rates similar to those of a previous face-to-face survey of drug use (66 percent to 72 percent). Drug use rates in the statewide telephone survey were similar to those found in the 1979 NIDA face-to-face household survey, thereby confirming experiments which had found no significant difference between telephone and face-to-face responses to a variety of interview questions (see Discussion). Compared to personal interviews, telephone interviews represent an attractive approach for localities with limited budgets since telephone in-

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Address reprint requests to William E. McAuliffe, PhD, Associate Professor, Department of Psychiatry, Harvard Medical School at Cambridge Hospital, 1493 Cambridge Street, Cambridge, MA 02139. Dr. Breer is Lecturer at that department; Dr. Ahmadifar is Research Associate at North Charles Foundation in Cambridge, MA: Dr. Spino is Assistant Professor of Biostatistics at Harvard School of Public Health. This paper, submitted to the Journal February 27, 1990, was revised and accepted for publication August 13, 1990.

terviews cost an average of half as much as face-to-face surveys using identical questions.[7]

Measures of drug *use,* however, have limited value in estimating the number of persons in the population who need treatment. Effective planning requires a measure of clinical levels of drug *abuse.* In 1981, Robins and her colleagues developed the Diagnostic Interview Schedule (DIS)[8] for epidemiological studies of mental disorders including alcohol and drug abuse. The DIS combines answers to questions on the frequency and consequences of use to form reliable scales of abuse and dependence that satisfy the American Psychiatric Association's formal (DSM-III) criteria for drug abuse and dependence.[9] Studies have demonstrated that the DIS can be used successfully in face-to-face household interviews,[10] and its use permits comparison with other populations already reported in the literature.[11] When DIS scores were compared with clinical diagnoses based on patient charts and DSM-III checklists, results showed moderate agreement (kappas for individual symptom scales ranged from −.01 to .63 with an average of .22 in one study[12] and .30 in another[13]).

Methods

The primary source of drug abuse data for this study was a telephone survey of Rhode Island households. The survey employed a proportional, stratified, multistage sample of persons age 12 or over living in households with telephones. The telephone numbers were identified by random digit dialing. A random sample of telephone numbers was drawn from each of the state's eight drug treatment catchment areas, with each area represented by a total proportionate to its share of the state population. Assuming a response rate similar to those found in previous New York State drug surveys (66 percent to 72 percent), we planned to use 7,000 of these numbers to reach the goal of 5,000 interviews. Interviewers asked to speak to the person in the household (12 years or older) with the most recent birthday; no substitutions were permitted.

The sample statistics were weighted[14] to produce representative parameter estimates for catchment areas and the state as a whole. Each respondent received an initial weight equal to the total number of persons age 12 or over in the household; the number of respondents in each catchment area was also weighted by the actual population size of the area. The weights were adjusted to correct for differences in response rate among catchment areas and multiplied by a constant to make the weighted sample size equal to the unweighted sample size of 5,176 completed interviews.

To increase community acceptance of the survey, all treatment and prevention agencies were notified about it and the Governor mentioned the study in his State of the State address. Several articles appeared in Rhode Island newspapers urging community cooperation.

We used the 1986 New York State drug survey[6] for basic drug use items and the Diagnostic Interview Schedule[8] for measures of abuse and dependence. We designed questions to measure *interest* in obtaining treatment. The final instrument was translated into Spanish for Spanish-speaking respondents.

Interviewers called each number in the sample until they succeeded in interviewing the designated respondent; determined that the number was either a business phone, had been disconnected or never assigned to a household by the telephone company; or reached the cut-off point of seven attempts. Interviewers made their first few calls in the evening (5:00 to 9:00); if these were not successful, they made further attempts on weekends. As a last resort, they tried calling on weekday mornings and afternoons. When a respondent refused to be interviewed, the survey supervisor arranged to have a more experienced interviewer try again several weeks later.

Reliability and Validity Studies

For 100 consecutive interviews conducted by several different interviewers, we compared the answers recorded by a supervisor listening on a remote monitor with the answers recorded by the interviewers. To estimate the potential bias of respondents who refused to cooperate, we compared the responses of persons (N = 130) who initially refused to respond but agreed to be interviewed when called again, to the responses of individuals who never refused. To estimate the potential bias of households that failed to answer the telephone, our research staff called all numbers not reached after seven attempts by the subcontractor. We then compared the responses of hard-to-reach individuals (the 92 reached by our staff) to those of respondents who were reached with seven or fewer calls. We also compared the sample demographics with census data.

To evaluate the truthfulness of respondents, we conducted a "randomized response" experiment in which a randomly selected subsample of respondents (N = 167) flipped three coins to determine whether or not to answer questions on drug use truthfully.[15] Since these respondents tossed the coins at home, interviewers could not tell whether the respondents were answering truthfully or just obeying the dictates of the toss ("yes" if all heads, "no" if all tails). Aggregate responses were statistically adjusted for expected frequencies. We then compared the resulting frequency of drug use reported by experimental subjects with the frequency of drug use reported by respondents in the general survey.

Drug Use and Abuse Measures

We estimated the percentage of lifetime and current moderate users, potential abusers and abusers. *Potential abusers* were respondents who reported "pathological" drug use (i.e. use of a drug daily for two weeks or more without a prescription) *or* that they had tried to cut down and failed to do so. *Abusers* were those who admitted to *both* pathological use *and* significant social impairment as a result of that use *or* to being drug dependent (showed signs of withdrawal and tolerance) whether or not they met the other two criteria.[8]

Computation of Unmet Demand

We defined the effective demand for drug treatment services as the number of services likely to be requested at current prices, assuming widespread knowledge of the availability of services and reasonable accessibility to them. We divided effective demand into *met demand* (the aggregate of clients enrolled in a drug treatment program at some time during the last twelve months) and *unmet demand* (the number of abusers who had not been in treatment during that period but who represented suitable candidates for admission). We asked all respondents who qualified as abusers by DIS criteria whether they were in treatment and, if not, whether they wanted treatment or at least had ever thought about entering treatment.

A conventional approach to estimating effective demand involves multiplying the estimated number of abusers in a population by a fixed percentage of abusers thought likely to seek treatment, the rule of thumb being 15 percent to 20 percent.[4] Using estimates of drug abuse prevalence from the telephone survey, we also estimated demand for treatment using this fixed percentage method.

TABLE 1—Validity Studies on Subpopulations

% Ever Used	Converted Refusals (n = 130)	Never Refused (n = 5046)	Difference (95% CI)	Hard to Reach (n = 92)	Not Hard To Reach (n = 5084)	Difference (95% CI)	Randomized Response* (n = 167)	Main Survey (n = 5176)	Difference (95% CI)
Marijuana	10.7	24.4	−13.7 (−19.2,−8.2)	29.7	24.0	5.7 (−3.7,+15.1)	28.0	24.1	3.9 (−3.0,+10.8)
Cocaine	3.0	7.3	−4.3 (−7.3,−1.3)	8.6	7.1	1.5 (−4.3,+7.3)	8.5	7.2	1.3 (−3.0,+5.6)
LSD/PCP	3.0	3.7	−0.7 (−3.7,+2.3)	3.0	3.7	−0.7 (−4.2,+2.8)	3.3	3.7	−0.4 (−3.2,+2.4)
Heroin	0.0	0.5	−0.5 (−0.7,−0.3)	0.0	0.5	−0.5 (−0.7,−0.3)	−7.5	0.5	

*Frequencies have been statistically adjusted for forced choice. The negative frequency for heroin use suggests that some respondents in the experiment could not bring themselves to say "yes" to having used heroin, even when the coin toss dictated that answer. The percent saying "yes" (6.9) was less than expected on the basis of chance alone. Subtracting the actual percent from the expected percent produced a negative number. Without additional data, that negative frequency is uninterpretable. Because the other three drugs showed no significant difference, it seems reasonable to conclude that the same was true for heroin.

Community Indicators of Drug Abuse

We analyzed established drug abuse indicators[16–18] to validate the telephone survey findings and to estimate optimal geographic distribution of services. Indicators were drug treatment admissions, hospital discharges (primary or secondary diagnoses of drug overdose, and so on), from the Rhode Island Department of Health, drug-related criminal complaints (suspected use, possession, or sales) from the Rhode Island Division of Drug Control, and drug arrests from the Federal Bureau of Investigation.

Allocating New Slots

We were asked to estimate the number of new treatment slots needed (if any), and to recommend how those slots should be distributed across both treatment modalities and catchment areas. To do so, we first looked at the drug use patterns of those respondents who said they wanted treatment now. Opiate users (who might be appropriate for methadone maintenance) were separated from users of other drugs. We then considered the current mix of treatment modalities in Rhode Island and used interviews with providers and state officials to identify gaps in modalities and levels of patient care. We also weighed the relative cost-effectiveness of different treatments. We estimated the optimal geographic distribution of admissions from their regression on survey drug abuse rates, criminal complaints, and hospital discharges.

Interviews with Providers and State Officials

A staff sociologist interviewed 38 drug program directors and state officials regarding current programs and problems involving client access, availability of services, AIDS, special needs, and other issues encountered in day-to-day operations, such as financing mechanisms, regulatory functions, relationships among agencies, and the continuum of care. To broaden our understanding of policy alternatives, we interviewed professionals from 13 states outside of Rhode Island.

The total cost (including indirect expenses) of the needs assessment was $150,000, of which our subcontractor received $50,000 ($10 per completed interview). The entire study (including a final report) took exactly six months to complete.

Results

Validity of Survey Findings

Of the 7,364 eligible households in the sample, 5,176 (70.3 percent) completed the interview. In the 100 consecutive cases where a supervisor coded responses independently of the interviewer, 99.3 percent of the responses were coded identically. Since most of the inconsistencies involved a question about friends' behavior, we excluded that item from further analysis.

Fifteen percent of those who initially refused to be interviewed eventually agreed to participate in the study. These 130 "converted refusals" reported *lower* frequencies of drug use and abuse than did those respondents who never refused (Table 1). The differences, however, were small except for ever-use of marijuana.

A comparison of 92 hard-to-reach respondents with those that were reached without extra effort revealed no significant differences in either drug use or abuse (Table 1).

The 167 respondents who were given the opportunity to disguise their answers in the "randomized response" experiment were no more likely to report illicit drug use than those in the general survey who were asked the same questions directly (Table 1).

A comparison of the survey sample with 1980 census figures for Rhode Island indicates close agreement on age, gender, and ethnicity.[19] A regression of catchment area drug abuse prevalence rates (measured by the survey) on drug-related hospital discharges and drug complaints provides further evidence ($R^2 = .89$) of the validity of the survey results.

Extent of Drug Use

Our survey yielded an estimated total of 13,948 current abusers, less than 2 percent of the state's 12 or older population of 793,334 (Table 2); 8,430 residents qualified as current potential abusers and 60,388 additional residents qualified as current minimal-to-moderate users. Approximately one-fourth of Rhode Island residents age 12 or older admitted to having *ever* used at least one illicit drug (marijuana, cocaine, LSD, or heroin) or licit psychoactive drug (sedative, amphetamine, or prescribed pain killer) without a prescription. The figure for *current* use (at least once in the past year) was 10.4 percent. Marijuana was used and abused more than any other drug, followed by cocaine, opiates, sedatives, amphetamines, and hallucinogens in that order.

Unmet Demand for Treatment

To determine unmet demand for treatment, we divided abusers not in treatment into a matrix of five levels of interest in treatment by three levels of the currency of abuse (Table 3). Analysis of the matrix generated 5,365 abusers (rows 1 and 2, and "recent" and "present abusers" in row 3) that made up the unmet demand for drug abuse treatment. We assumed that "past abusers" and "recent abusers" in row 1 were using currently

TABLE 2—Estimated Current Users, Potential Abusers, and Abusers, by Type of Drug

In Last Year	At Least One	Marijuana	Cocaine	Opiates**	Sedatives*	Amphetamines*	LSD/PCP
Users (Minimal or Moderate)	60,388	49,813	18,239	10,116	6,591	5,364	3,525
Potential Abusers	8,430	7,664	1,839	153	153	1,073	0
Abusers	13,948	10,269	4,292	1,839	2,605	1,839	766
Total	82,766	67,746	24,370	12,108	9,349	8,276	4,291
(% of R.I. pop.)	(10.4)	(8.5)	(3.1)	(1.5)	(1.1)	(1.0)	(0.5)

*Used on own (i.e., without prescription).

**Includes heroin, and pain killers used without a prescription.

TABLE 3—Levels of Unmet Demand for Treatment

Level	Definition	Past Abusers[a] (33,414)	Recent Abusers[b] (6,284)	Present Abusers[c] (7,089)
1	Want Treatment? Yes	307	920	920
2	Want Treatment? Not sure	460	0	460
3	Ex-clients[d]	2,146	1,073	1,225
4	Thought about Treatment[e]	1,686	460	460
5	Never Thought about Treatment[f]	28,815	3,831	4,024

[a] Have abused in past but not in last year.

[b] Have abused in past year but not in last month.

[c] Have abused in last month.

[d] Former drug clients who said "no" to wanting treatment now.

[e] Said "no" to wanting treatment now, have never been in treatment, but have considered treatment in past.

[f] Said "no" to wanting treatment now, have never been in treatment, and have never considered treatment.

but not at the point of abuse, or were in early recovery. The "not sures" in row 2 were included on the assumption that they would decide to enter a program if openings at the most sought-after services (methadone maintenance and short-term residential programs) were not in such short supply. Despite not wanting treatment now, the "present abusers" in row 3 were included because they had sought treatment before, had relapsed, but, we believed, would want treatment again were it readily available. The "recent abusers" in row 3 would similarly want intermediate care or aftercare. We excluded the respondents in row 3 who had not abused for over a year on the assumption that they were sufficiently recovered.

To get the total demand for treatment, we added unmet demand and met demand. Rather than estimating met demand from the number of survey respondents *currently* in treatment (which would exclude individuals discharged earlier in the year), we used actual MIS admissions (2,212) for the previous year. In this way, we estimated total effective demand to be 7,577 in 1986. We recommended that Rhode Island should expand its treatment system until it could accommodate all 7,577.

According to the traditional rule of thumb for estimating the number of addicts who will avail themselves of treatment at any one time, there were 2,789 Rhode Island residents (20 percent of the estimated 13,948 current drug abusers in the State) who were candidates for drug abuse treatment at the time of our study. This estimate suggests that if Rhode Island were to provide treatment for all addicts who wanted it, it would have to expand its services by 26 percent (from 2,212 to 2,789) as opposed to the 243 percent expansion (2,212 to 7,577) indicated by the first method.

Expansion of Services

Given the State's fiscal constraints and siting difficulties for new facilities, we recommended that the expansion of services be divided into two phases. In phase I (two to three years) the State would add enough slots to accommodate the estimated 2,150 abusers who said in the telephone interview that they wanted treatment right now (all abusers from row 1). Adding 2,150 admissions would nearly double the present treatment system. In phase II, which would extend for several years beyond phase I, the State would add enough slots to accommodate the remaining 3,200 abusers (all abusers from row 2 and most from row 3) whom we deemed likely applicants for new treatment slots but whose desire to enter a program immediately was less definite under present conditions. A phased expansion would allow for adjustments if drug abuse indicators suggested substantial changes in addict prevalence during phase I or if our conclusions about demand for services were not borne out by experience.

For most treatment modalities, we recommended making phase I increases proportionate to the number of slots already in existence. We assumed that the mean number of admissions per slot would remain constant in the future and that the current mix of services reasonably reflected the demand for different types of services. We suggested a somewhat larger expansion for methadone maintenance programs, however, because half of the abusers who said they wanted treatment reported heroin addiction, and because, at the time of the study, there was a two-year waiting period at the State's only methadone center. We also recommended adding a small number of residential 28-day drug treatment slots and a still smaller number of outpatient narcotic antagonist (Naltrexone) slots[20,21] to serve the opiate addicts. Neither of these latter two modalities formally existed in Rhode Island at the time of the study.

To guide geographic distribution of new treatment slots, we regressed treatment admissions on a weighted index of drug-related hospital discharges, drug-related complaints to the Division of Drug Control, and our survey results on the number of drug abusers in each catchment area. For outpatient drug-free admissions (including day treatment), the regression of treatment capacity on the weighted index of hospital discharges, drug complaints, and abusers revealed in the survey yielded an R^2 of .55. The results (Figure 1) confirmed the opinion expressed by some officials that treatment facilities were overly concentrated in central regions of the state. We recommended that proportionately more new admissions should be added according to the discrepancy between actual and expected number of slots.

To determine where new methadone slots should be sited, we performed a similar regression analysis and comparison using opiate admissions as the dependent variable and opiate use as reported in the survey as the independent variable. In the case of short-term (28 day) and long-term (therapeutic community) residential modalities where the number of existing slots was too small and concentrated for a regression analysis, we adopted the rule of thumb that the State should locate new slots in areas that presently had no residential facilities for drug abusers. This made sense in light of the survey finding that there were significant numbers of drug abusers in every catchment area.

Interview Results

Providers and state officials were in agreement that the primary problem was the inadequacy of treatment capacity. While providers stressed shortage of staff and low salaries, state officials stressed the need for mandation of third-party reimbursement of substance abuse services, quality assurance, and greater access to treatment for subpopulations such as adolescents and the poor. We used these interview results in recommending increased funding and new mechanisms for improving utilization of existing funds (revision of traditional lengths of stay and establishment of a statewide panel to review treatment practices). Interviews with policymakers in other states provided us with valuable information on some of the problems to be expected with third-party mandation such as changes in the number of drug abusers seeking care and increases in premium cost. We also recommended for-profit programs and "medical maintenance"[22] for increasing methadone slots in the face of limited public allocations.

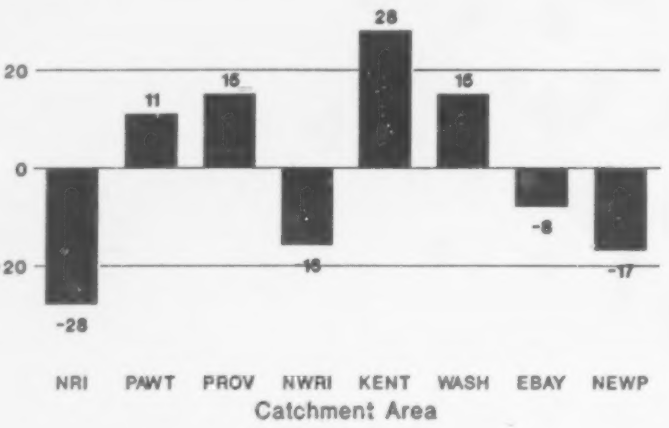

FIGURE 1—Adequacy of treatment allocation: Actual minus expected number of OPDF slots,* by catchment area.

*Estimated by statistically regressing OPDF slots on an index made up of: 1) current abuse rate of *any* drug; 2) number of drug-related hospital discharges; and 3) number of complaints reported to the RI Division of Drug Control (OPDF Slots = .67 + .07 (Drug Problem Index).

Reaction to Recommendations

The initial reaction to our report was mixed. In a series of hearings and written statements, most state officials praised it.[19] Third-party payors supported the report's recommendations for reduced lengths of stay and a higher ratio of outpatient to inpatient treatment, while objecting to its endorsement of mandatory drug treatment benefits (several months before we submitted our final report, the Rhode Island legislature had mandated third-party coverage for drug abuse treatment). Providers were "particularly impressed with the depth of the household survey and the methodology used to determine effective demand for treatment,"[19] but felt that we had paid too much attention to cost containment and quality regulation.

The Division of Substance Abuse, its parent organization, the Department of Mental Health, Retardation and Hospitals, and the Department of Health endorsed our recommendation that the drug treatment system be doubled in the next few years. In the year following submission of our report (1988), the State provided $700,000 for new methadone maintenance slots and another $3,900,000 for the construction and operation of a 60-bed adolescent treatment facility which our report had recommended. Private agencies added 300 methadone slots, 600 outpatient drug-free slots, and 60 28-day residential treatment slots. With the exception of the $700,000 for methadone maintenance and the projected adolescent facility, all of the expansion was funded by the private sector. By July 1989, the State's drug treatment network had expanded to levels close to those recommended for phase I of the expansion.[23] Methadone slots had reached 75 percent of the phase I target with a concomitant decrease in waiting lists. Outpatient drug-free slots had exceeded the phase I target and were already approaching the level we had recommended for phase II. The number of inpatient 28-day slots had already surpassed the phase II target.

The only major drug treatment modality to lag behind in the expansion was long-term residential care (therapeutic communities), the primary form of treatment excluded from the 1987 bill making third-party insurance coverage of drug abuse treatment mandatory throughout Rhode Island. In our follow-up report integrating findings from the Brown (alcohol) and Harvard (drugs) reports, we made a special plea for legislation to correct this omission.*

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*McAuliffe WE, Lewis D, Breer P, Williams C: *Rhode Island integrated alcohol and drug treatment program matrix*. Cranston, Rhode Island: Division of Substance Abuse, Rhode Island Department of Mental Health, Retardation and Hospitals, unpublished report.

Throughout the expansion of the treatment network, the State exerted a continuing influence through its control over the certificate of need (CON) process. Once the Department of Health and related departments approved our recommendations, the State Health Planning Committee formally adopted the report's treatment slot recommendations as a guide in evaluating applications for new treatment facilities (both private and state funded).

Discussion

The percent (25) of Rhode Island residents age 12 or older who reported using illegal drugs at some point in their lives agrees with the results of national studies[24,25] and studies of nearby states.[6] Estimating the effective demand for drug treatment from survey results produced a figure almost three times as large as the total estimated by the fixed percentage method which assumes that only 20 percent of all abusers will seek treatment, the major difference stemming from our discovery of past abusers who want treatment and former clients who are currently abusing. The validity of the larger estimate is supported by the fact that Rhode Island doubled its treatment slots in the two years following submission of our report[23] and has thus far reported no difficulty filling those new slots.

The validity of estimating drug use and abuse by interview surveys has been studied since the 1960s. Measures of drug-use frequency have shown high internal and test-retest consistency among both addicts and non-addicts.[26,27] Surveys of special populations (addicts in treatment,[26,28–32] psychoactive prescription drug users,[33] Vietnam military returnees,[34] and addicts once known to the police[35]) have successfully corroborated interview statements of drug use with official records, family members, treatment personnel, and urinalysis results. Arrestees awaiting trial[36] are one group known to underreport substantially their use of illicit drugs.

There is both direct and indirect evidence that drug use survey responses in general populations are also reasonably valid.[37] The most direct evidence comes from use of the "randomized response" technique.[38] Respondents who are given a chance to conceal their true answers from the interviewer have generally reported levels of drug use equal to or only slightly higher than those of respondents asked to state their answers openly.[15,39] Indirect evidence can be found in the similarity of drug-use rates reported by different investigators studying independent but demographically comparable segments of the population.[37] Further support for the validity of drug-use questionnaires comes from the finding that, when respondents are asked about fictitious drugs, only a small fraction report any use.[40–42]

Studies comparing telephone survey responses with those of face-to-face interviews have generally found the results to be similar.[43,44] A recent study[45] found, however, that compared to face-to-face interviewing, telephone interviewing yielded significantly lower estimates of marijuana use among Blacks. The differential persisted after controlling for demographic differences between the two samples and excluding non-telephone households, suggesting that it had something to do with the mode of interviewing itself. However, the difference in rates did not hold for Whites, nor did it hold for cocaine use among either Blacks or Whites. Other controlled drugs were not tested. If there is a biasing effect in telephone interviewing, it is very limited and unlikely to affect overall prevalence rates of drug use and abuse in general surveys, particularly in surveys of states like Rhode Island with small minority populations.

Thus, research on drug surveys in general and telephone interviewing in particular suggests that under ordinary circumstances respondents tell the truth about their drug use. Where systematic bias occurs, it is typically in the direction of underreporting the frequency of drug use.[37,42] It follows that a treatment needs assessment based on community prevalence estimates of drug use and abuse tells us the *minimum* number of slots needed if the community is to make treatment available to all who need and would seek it. As for our own survey, several indicators discussed earlier suggested that the findings have acceptable reliability and validity.

One of the drawbacks of the household survey is that it yields proportionately few individuals who qualify as drug abusers. In our sample, less than two percent met criteria for drug abuse, and less than three-tenths of 1 percent expressed interest in getting treatment. To enlarge the base of abusers in order to investigate detailed planning issues, we suggest employing a larger total sample or, where cost considerations rule that option out, oversampling geographic areas known by arrests and hospital data to contain many drug addicts.

A drawback specific to the telephone drug abuse survey is that it excludes a subpopulation (households without a telephone) likely to have a high prevalence of drug abuse. This population, however, is quite small. In 1981, 98 percent of Rhode Island households had telephones.[46] Intravenous drug users (IVDUs) are a subgroup of drug abusers that one might expect to have the lowest telephone rates of all. In nearby Massachusetts, unpublished data from our current multi-city study of IVDUs (N = 577) not in treatment who were recruited through street outreach, indicates that 84 percent either have their own phone (38 percent) or have access to a phone at their current residence (46 percent).

Another limitation of any household survey, whether telephone or face-to-face, is that it excludes both the homeless and the institutionalized (prisoners, psychiatric patients, college dorm residents). The prevalence of drug abuse is known to be higher in these populations than in the general household population.[47] New York State's Division of Substance Abuse has recently developed a strategy for surveying the homeless, utilizing prearranged telephone interviews with randomly chosen residents of urban shelters.[47] In many states, drug treatment services for prisoners and psychiatric patients are provided by agencies other than the drug treatment agency and therefore may not be included in a needs assessment.

A recent US Senate report recommended improving upon survey estimates of cocaine abuse prevalence by adding nonsurvey data on nonhousehold populations (arrestees, drug patients, and the homeless).[48] Although we believe that the report's methodology may over-correct survey estimates, the report actually estimated *fewer* cocaine addicts in Rhode Island (3,900) than we estimated from our telephone survey (4,300).[19]

The orderly expansion of treatment services in Rhode Island testifies to the effectiveness of the needs estimation methods used in the present study. With improvements in survey methodology, statistical techniques, and the development of standardized diagnostic instruments suitable for testing large populations, substance abuse planning technology has entered a new era. It is reasonable to expect further progress in years to come. □

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Racial/Ethnic Differences in Smoking, Drinking, and Illicit Drug Use among American High School Seniors, 1976–89

Jerald G. Bachman, PhD, John M. Wallace, Jr., MA, Patrick M. O'Malley, PhD, Lloyd D. Johnston, PhD, Candace L. Kurth, MPH, and Harold W. Neighbors, PhD

ABSTRACT

Background. This paper reports racial/ethnic differences in the use of licit and illicit drugs by high school seniors in the United States.

Methods. The study uses questionnaire data from annual, nationally representative surveys of seniors from 1976 through 1989. Combined sample sizes were 57,620 for 1976–79; 75,772 for 1980–84; and 73,527 for 1985–89.

Results. Native Americans had the highest prevalence rates for cigarettes, alcohol, and most illicit drugs; White students had the next highest rates for most drugs. Asian Americans had the lowest prevalence rates, and Black students had levels nearly as low except for marijuana. Prevalence rates for the Hispanic groups were mostly in the intermediate ranges except for relatively high cocaine use among the males. Trend patterns for most forms of drug use were similar across subgroups, although cigarette use declined more sharply for Black than White seniors, resulting in greater Black-White differences in recent years.

Conclusions. This study, other school-based studies, and general population surveys all show relatively low levels of drug use by most non-White youth, especially Black Americans and Asian Americans. Multivariate analyses indicate that such subgroup differences in high school seniors' drug use are not primarily attributable to family composition, parents' education, region, or urban-rural distinctions. (*Am J Public Health* 1991;81:372–377)

Introduction

Drug use among American youth continues to be a focus of attention for politicians, educators, the media, and the general public. Although the media frequently focus on large cities, the nation's drug problems are not limited to any one group or to any one geographic area. The impacts of drug use and abuse are felt more keenly in the Black, Hispanic, and Native American communities,[1] but the patterns of drug use by members of these groups, particularly adolescents, are not well documented. The purposes of the present report are: to document the extent to which drug use varies among racial/ethnic subgroups of high school seniors; to explore whether subgroup differences have been changing in recent years; and to consider some of the implications of such differences.

The data reported here are based on samples of high school seniors, representative of the United States as a whole. The study does not undertake to explore substantial differences in drug use which may exist from one city to another, or from one neighborhood to another. More importantly, surveys of high school seniors do not include those youth who drop out before graduation; thus, although this report represents most young people, including the majority of those in each racial/ethnic subgroup studied, it does not cover the total cohort of 17- to 18-year-olds. These limitations are treated at some length in the discussion.

Methods

The Monitoring the Future project, conducted by the University of Michigan's Institute for Social Research under grants from the National Institute on Drug Abuse, has surveyed large, nationally representative samples of high school seniors during the spring of each year since 1975. The design and procedures are summarized briefly below; detailed descriptions are presented elsewhere.[2-4] The measures of drug use and racial/ethnic identification are presented in the Appendix.

Samples

A three-stage sampling procedure is employed,[5] with stage 1 the selection of particular geographic areas, stage 2 the selection of one or more high schools in each area, and stage 3 the selection of seniors within each high school. The result each year is an area probability sample of the 48 coterminous states. About 130 high schools participate each year (approximately 118 public and 12 private). About 83 percent of the sampled seniors generally participate (nearly all nonparticipation results from absenteeism), yielding samples of about 17,000 seniors each year.

Survey Procedures

Data are collected via questionnaires administered in classrooms by locally based Institute for Social Research representatives and their assistants, following carefully standardized procedures. The questionnaires are designed for optical scanning; all items are closed-ended. Five different questionnaire forms have been

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From the Institute for Social Research, University of Michigan. Address reprint requests to Jerald G. Bachman, PhD, Institute for Social Research, University of Michigan, Ann Arbor, MI 48106-1248. This paper, submitted to the Journal April 27, 1990, was revised and accepted for publication September 5, 1990.

TABLE 1—Racial/Ethnic Distribution of the Monitoring the Future Combined Samples*

	1976–79		1980–84		1985–89	
	N	%	N	%	N	%
White	48993	85.0	62157	82.0	57864	78.7
Black	5965	10.4	8958	11.8	8187	11.1
Mexican American	1088	1.9	1480	2.0	3117	4.2
Puerto Rican/Latin American	572	1.0	1105	1.5	1392	1.9
Asian American	439	0.7	1139	1.5	1899	2.6
Native American	563	1.0	933	1.2	1068	1.5
Total	57620	100.0	75772	100.0	73527	100.0

*The multistage sampling design with respondents clustered in schools produces larger sampling errors than would a simple random sample of equivalent size. For statistics in the present paper, the estimated design effects are 8.0 for White males and females, 3.0 for Black males and females, and 2.2 for the males and females in any of the other groups. Frequencies used to calculate statistical significance are equal to the actual number of cases shown in the table divided by the appropriate design effect (i.e., 8.0, 3.0, or 2.2). When referring to group differences the term "significant" or "significantly" refers to $p < .01$.

used each year, with each administered to a random one-fifth of the sample.

Statistical Significance and Confidence Intervals

Confidence intervals vary greatly depending upon sample size, design effects (see note to Table 1), and percentage size. As examples and general guidelines, we note the following: The *largest* 95% confidence intervals around percentages in Tables 2–4 are 1.7 percent for Whites, 2.8 for Blacks, 3.7 for Mexican Americans, and 6.3 for Native Americans (the smallest subgroup). Any Black-White difference equal to or exceeding 3.0 percent is significant at the 95% confidence level. It would be unnecessarily cumbersome to specify significance levels for every subgroup comparison discussed herein; instead, we have adopted the convention of describing *as differences* only those which exceed the 99% confidence level.

Results

Our primary emphasis is upon racial/ethnic subgroup differences in drug use, especially those which have been somewhat consistent across time. We examined annual data from the senior classes of 1976 through 1989, combining the classes into three groups: 1976–79, 1980–84, and 1985–89. The resulting numbers of cases for the total samples and the racial/ethnic subgroups are displayed in Table 1.

Table 1 indicates that the proportions of Whites in the samples decreased over time, whereas the proportions of Asian Americans and Hispanics increased substantially. These changes generally parallel shifts in the overall population of youth, although it should be kept in mind that the proportions in our samples reflect some substantial subgroup differences in high school dropout rates, as well as other differences (including a somewhat higher rate of missing data for several of the subgroups other than White). In particular, it should be noted that nationwide dropout rates for Blacks have declined in recent years, and are now only slightly higher than those for Whites.[6]

Subgroup Differences in Prevalence of Drug Use

Marijuana

Annual prevalence rates for marijuana (1985–89 combined; see Table 2) are highest among Native American females and males, and nearly as high among White males and females, and Mexican American males; rates are somewhat lower for Mexican American females, Puerto Rican and other Latin American males, and Black males; and rates are lowest among Puerto Rican and other Latin American females, Black females, and Asian American males and females. Monthly and daily (or nearly daily) prevalence rates for marijuana show similar subgroup distinctions, except that male-female differences are more pronounced (Tables 3 and 4).

Cocaine

Prevalence rates for cocaine are highest for Native Americans (males and females) and males in both Hispanic groups, but significantly lower for Hispanic fe-

TABLE 2—Annual Prevalence of 13 Types of Drugs, 1985–89 Data Combined, by Sex and Race

	Percent who used in last 12 months											
Types of Drugs	White Male	Black Male	MexAm Male	PR&LA Male	Asian Male	Nat Am Male	White Female	Black Female	MexAm Female	PR&LA Female	Asian Female	Nat Am Female
Minimum N =	(28056)	(3688)	(1518)	(680)	(982)	(537)	(29808)	(4499)	(1599)	(712)	(917)	(531)
Marijuana/Hashish	40.2	29.8	37.3	30.6	19.6	42.0	36.0	18.4	26.0	21.3	17.1	44.0
Inhalants[a]	8.8	2.6	6.0	5.1	4.8	9.6	5.2	2.2	4.3	2.9	3.2	4.4
Hallucinogens	8.3	1.9	5.9	6.5	3.0	10.0	5.0	0.6	2.2	2.1	2.2	9.0
LSD	7.0	1.3	5.2	3.4	2.5	7.8	3.9	0.3	1.6	1.1	1.9	7.2
Cocaine	11.9	6.1	14.7	15.6	5.8	14.2	9.3	2.6	7.6	8.2	5.7	15.5
Heroin	0.7	0.7	0.9	1.2	0.4	1.5	0.3	0.4	0.4	0.4	0.2	1.0
Other opiates[b]	6.5	1.9	3.2	3.0	3.1	7.4	5.3	1.2	2.1	1.6	2.1	5.7
Stimulants[b]	13.6	4.6	11.3	8.0	5.6	17.0	14.7	3.1	10.1	5.9	7.0	19.4
Sedatives[b]	5.3	2.2	4.7	4.6	3.4	8.8	4.4	1.2	2.7	2.6	2.6	6.4
Barbiturates[b]	4.4	1.9	4.1	4.0	2.6	7.2	3.8	1.1	2.4	2.5	2.3	6.2
Methaqualone[b]	2.5	0.9	1.2	2.3	1.5	4.8	1.4	0.3	0.5	0.5	0.9	2.2
Tranquilizers[b]	5.8	1.7	2.6	3.1	3.2	6.9	5.9	1.4	2.1	4.1	1.8	8.7
Alcohol	88.3	72.5	82.4	80.6	69.3	82.0	88.6	63.9	73.6	77.2	67.5	81.3

[a] Data based on four questionnaire forms. N is four-fifths of N indicated.
[b] Only drug use which was not under a doctor's orders is included here.

TABLE 3—Thirty-Day Prevalence of 14 Types of Drugs, 1985–1989 Data Combined, by Sex and Race

	Percent who used in last 30 days											
Types of Drugs	White Male	Black Male	MexAm Male	PR&LA Male	Asian Male	Nat Am Male	White Female	Black Female	MexAm Female	PR&LA Female	Asian Female	Nat Am Female
Minimum N =	(28056)	(3688)	(1518)	(680)	(982)	(537)	(29808)	(4499)	(1599)	(712)	(917)	(531)
Marijuana/Hashish	25.0	18.5	22.0	18.9	9.7	27.6	19.8	9.9	13.6	9.6	8.1	23.9
Inhalants[a]	3.4	1.4	2.3	2.0	1.3	5.2	2.0	1.4	2.1	0.8	0.8	0.9
Hallucinogens	3.5	0.9	2.4	3.0	1.5	3.6	1.7	0.3	0.7	0.4	0.3	2.7
LSD	2.8	0.6	1.9	1.6	1.1	3.1	1.1	0.2	0.3	0.2	0.1	2.2
Cocaine	5.6	2.6	8.2	8.1	1.8	7.3	4.1	1.3	3.0	2.9	2.6	9.2
Heroin	0.3	0.5	0.3	0.9	0.1	1.1	0.1	0.3	0.2	0.2	0.0	0.4
Other opiates[b]	2.3	0.9	1.1	1.5	1.6	4.0	1.9	0.6	0.7	0.5	0.7	2.4
Stimulants[b]	5.6	1.9	4.9	3.1	2.1	8.1	6.0	1.3	4.8	1.2	3.6	10.3
Sedatives[b]	2.2	1.1	2.0	1.8	1.9	4.8	1.7	0.5	0.9	1.3	1.3	2.6
Barbiturates[b]	1.8	0.9	1.7	1.3	1.4	3.7	1.5	0.5	0.8	1.2	1.0	2.1
Methaqualone[b]	0.9	0.5	0.6	0.9	0.8	2.5	0.5	0.1	0.2	0.1	0.6	0.9
Tranquilizers[b]	1.9	0.8	0.8	0.6	1.7	3.1	2.0	0.5	0.9	1.5	0.9	2.2
Alcohol	72.3	49.2	65.0	55.4	43.7	69.0	66.6	32.8	50.5	43.0	34.2	60.2
Cigarettes	29.8	15.6	23.8	22.0	16.8	36.8	34.0	13.3	18.7	24.7	14.3	43.6

[a] Data based on four questionnaire forms. N is four-fifths of N indicated.
[b] Only drug use which was not under a doctor's orders is included here.

TABLE 4—Daily Use of Three Types of Drugs in the Last 30 Days, 1985–1989 Data Combined, by Sex and Race

	Percent who used daily in last 30 days											
Types of Drugs	White Male	Black Male	MexAm Male	PR&LA Male	Asian Male	Nat Am Male	White Female	Black Female	MexAm Female	PR&LA Female	Asian Female	Nat Am Female
Minimum N =	(28056)	(3688)	(1518)	(680)	(982)	(537)	(29808)	(4499)	(1599)	(712)	(917)	(531)
Marijuana/Hashish	5.1	2.6	4.2	3.5	1.7	8.2	2.1	0.9	1.1	0.5	0.5	4.3
Alcohol												
Daily	7.0	4.2	8.3	4.0	2.3	10.1	2.8	0.7	2.6	0.9	0.9	5.4
5+ drinks in a row/last 2 weeks	48.1	24.0	45.3	31.4	19.4	48.1	31.3	9.3	23.6	14.5	10.7	33.7
Cigarettes	18.8	8.6	11.6	13.3	9.0	26.0	22.5	7.1	8.1	13.3	9.4	33.8
Half-pack or more per day	12.5	3.3	5.2	6.1	4.4	18.4	13.3	2.2	2.5	4.2	4.5	23.4

males; rates for Whites are fairly similar to those for Hispanics, except that the difference is somewhat smaller between males and females; and rates are significantly lower for Blacks and Asian Americans. In the case of Asian Americans, prevalence rates for males and females are essentially the same; however, for Blacks, like Hispanics, the prevalence is about twice as high for males as for females.

Other Illicit Drugs

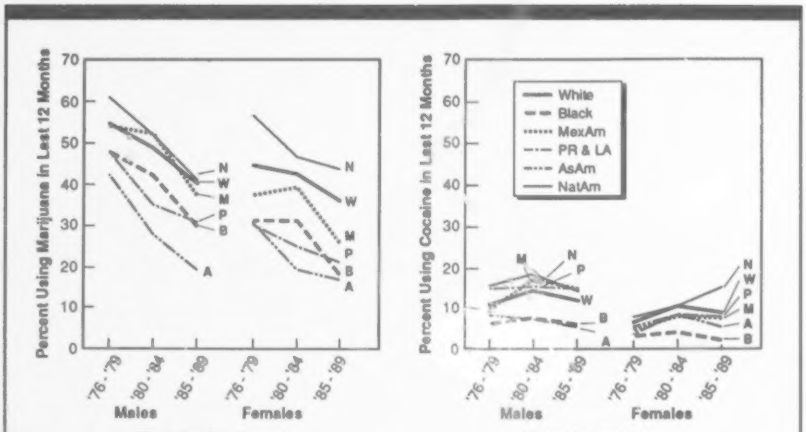
Tables 2 and 3 provide details on a number of other illicit drugs: inhalants, hallucinogens, heroin, other opiates, stimulants, sedatives, and tranquilizers (any use under doctor's orders is excluded). For most drugs and most subgroups, fewer females than males report use, with usage rates being highest for Native American seniors, and lowest for Black and Asian American seniors. Differences among White and Hispanic seniors are generally smaller and not consistent across these other illicit drugs.

Alcohol

Alcohol use among White and Native American males and females is relatively high, while among Black and Asian American seniors, only about half of the males and one-third of the females report use of alcohol during the past month. Substantial subgroup differences exist in frequent or heavy use of alcohol (i.e., five or more drinks in a single sitting). As shown in Table 4, almost half of White, Native American, and Mexican American males report such heavy alcohol use once or more during the two weeks preceding the survey. Heavy drinking is significantly less prevalent among Puerto Rican and other Latin American males, and even lower among Black males and Asian American males. Racial/ethnic differences among females generally parallel those for males, but at distinctly lower prevalence levels. Although relatively few seniors drink alcohol on a daily or near daily basis, this behavior shows subgroup differences which parallel those for heavy drinking (Table 4).

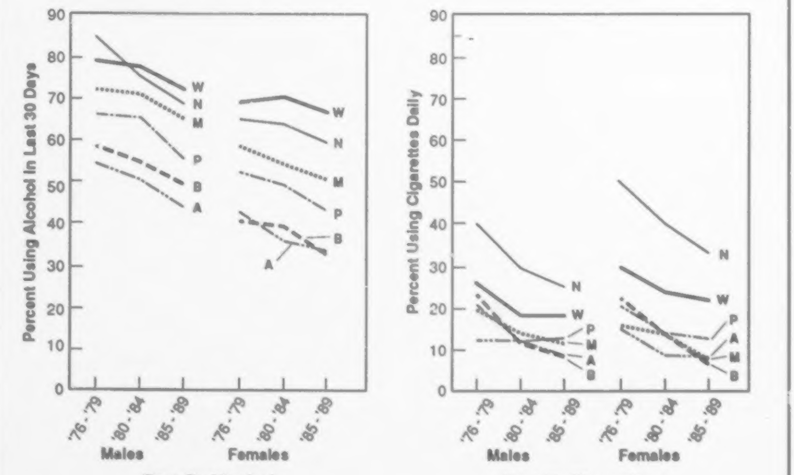
Cigarettes

About two-thirds of all seniors have tried cigarettes sometime in their lives;[2] however, the much more important measures are those reflecting their current behavior: monthly prevalence (Table 3), and especially daily prevalence and half-pack-a-day use (Table 4). Half-pack daily use is highest among Native American seniors, significantly lower among Whites, and far lower among the other subgroups. In contrast to most other drugs, cigarettes have been used just as much by young women as by young men in recent years.



Part A: Marijuana

Part B: Cocaine



Part C: Alcohol

Part D: Cigarettes

FIGURE 1—Trends in Use of Four Drugs, 1976–89, by Sex and Race

Trends in Prevalence of Drug Use, 1976–89

There have been a number of important increases and decreases in the use of various drugs since the mid-1970s, as we have reported in detail elsewhere.[2,7] Our purpose here is to consider whether some of the most important racial/ethnic subgroup distinctions observed in the classes of 1985–89 are relatively long-standing ones or reflect recent changes.

Illicit Drugs. Part A of Figure 1 shows that the downward trend in annual marijuana use occurred within all subgroups, and that subgroup differences were fairly consistent across time. Part B of Figure 1 shows the overall rise in cocaine use between the late 1970s and the early 1980s; the patterns of subgroup differences are not so consistent across time, but because of the generally small proportions of cocaine users we cannot be very confident in asserting different subgroup trends across time. Use of most of the other illicit drugs shown in Tables 2 and 3 declined during the past decade[2]; here again, the generally low proportions of users provide little basis for asserting *differential* subgroup trends across time. (Data for the earlier time periods are presented elsewhere.[8])

Alcohol. As Part 3 of Figure 1 indicates, 30-day prevalence of alcohol use has declined somewhat, especially among males, but subgroup differences were largely consistent across time.

Cigarettes. As shown in Part D of Figure 1, prevalence rates for daily use of cigarettes have declined for all subgroups. Substantial declines occurred between 1976–79 and 1980–84, whereas declines in recent years have been much smaller. Some groups (e.g., White males) showed no decline in recent years. In general, the declines in smoking have been stronger among non-White groups. Most notably, daily smoking prevalence rates of Black seniors dropped by two-thirds (from 23.6 percent to 8.6 percent for males, and 22.3 percent to 7.1 percent for females). Thus, Black-White differences in cigarette smoking have become more pronounced in recent classes of high school seniors.

Discussion

Although relatively little research compares drug use among racial/ethnic subgroups, the existing studies[7,9–20] yield a fairly consistent set of findings: Asian American youth tend to report very low levels of drug use relative to other groups. Black youth consistently report lower rates of drug use than White youth. Hispanic youth typically report more substance use than Blacks and only slightly less than Whites. Native American youth on average report greater use of alcohol and other drugs than any other subgroup.

The present findings are largely consistent with the literature. One contribution of the present research has been to document these differences based on large, nationally representative samples of high school seniors. Another contribution has been to demonstrate that for more than a decade the racial/ethnic subgroups have shown parallel trends, for the most part, in their use of alcohol and the illicit drugs. An important divergence in smoking rates (smaller declines among Whites) also has been noted.

Several important questions remain. First, are these replicated (and thus *reliable*) findings of subgroup differences also *valid*? In other words, how accurate are young people's self-reports of drug use in general, and are there reasons to suppose that subgroups may differ in accuracy? Second, how do the subgroups differ in dropout rates, and how do these differences affect our ability to generalize to total age cohorts? These are issues which have concerned us for some time, and they have made us cautious about reporting subgroup differences in drug use.[4] Although we have not fully resolved these questions, we have become increasingly confident that the subgroup differences reported here are, on the whole, valid.

First, with respect to self-reports of drug use, there is a growing literature indicating their validity.[21–25] Some questions may remain, of course, about the willingness of seniors to report their drug use honestly if they are not confident about the purposes of the research, if they do not trust the survey administrators, and/or if they have greater than average tendencies toward favorable self-presentation. Mensch and Kandel report that in the National Longitudinal Survey of Youth, which employed face-to-face interviews, Black and Hispanic youth were more likely than White youth to underreport their use of marijuana; however, this underreporting occurred most often at the lowest levels of use.[26] Incidentally, their initial analyses (based on total samples) revealed no underreporting of the licit drugs (i.e., alcohol and cigarettes), which show the largest subgroup differences in the present study.

Earlier analyses of Monitoring the Future data[8] revealed subgroup differences in rates of missing data and inconsistent responses; however, these differences do not parallel the subgroup differences in self-reported drug use. For example, Native American seniors have relatively high rates of missing data and/or inconsistent responses (6.6 percent for marijuana) as well as high self-reported drug use, whereas Black seniors have equally high rates of missing data and/or inconsistent responses (6.8 percent for marijuana), but much lower self-reported drug use. White and Asian American seniors are fairly similar in having low rates of missing data and/or inconsistent responses (2.3 percent and 3.1 percent, respectively, for marijuana), but their rates of drug use are distinctly different.

Additional analyses of Monitoring the Future data have shown that Blacks are more likely than Whites to perceive that drug use involves high risks, and to disapprove of drug use. Black seniors also are much less likely to report smoking, alcohol use, and drunkenness by friends, consistent with the large Black-White differences in self-reported drug use.* It is hard to imagine that these complexly interrelated findings are simply the result of selective distortion; in particular, it is unlikely that such distortion would result in much larger differences in self-reported use of the licit drugs compared with the illicit drugs. We find it more parsimonious to conclude that the substantial Black-White differences in self-reports are largely the result of genuine differences in drug use between our samples of Black and White high school seniors.

Even if self-reports are mostly valid across all of the subgroups studied, we are faced with further questions because our surveys do not include young people who drop out of high school before late spring of their senior year. Given that drug use is generally higher among dropouts, how different would findings be if based on the total cohort of 17- to 18-year-olds rather than just high school seniors? It is highly likely that the inclusion of dropouts would tend to raise the observed prevalence rates for all drugs and all subgroups. In addition, it seems fairly likely that the impact would be greatest in subgroups with high dropout rates, thus changing subgroup comparisons in several respects. First, Hispanics have higher than average dropout rates;[6] therefore, comparisons of *all* White and Hispanic 17- to 18-year-olds would probably yield somewhat smaller differences in drug use than found in samples of seniors. Second, Asian Americans have lower than average dropout rates[6]; accordingly, this subgroup might be even farther below average given drug use data based on the total age cohort rather than just seniors. Third, given the very high dropout rates among Native Americans,[6] total age cohort data might place them even farther above average in drug use. Fourth, Black-White differences in drug use might be reduced if we compared the total age cohort rather than just seniors; however, because Black and White dropout rates are now fairly similar,[6] the reduction would not be very large, unless drug use is differentially correlated with dropping out. It is also worth noting that recent household surveys, which do not omit dropouts, find Black-White differences in youthful drug use roughly as large as those reported here.[7]

If drug use really is lower than average among most non-White youth, an interesting question remains as to why this occurs. Multivariate analyses reported elsewhere[8] indicate that it cannot be attributed primarily to differences in parental presence or education, nor does it have a great deal to do with where seniors live throughout the United States. Some other dimensions of lifestyle we have measured have more bearing on drug use, but they do not account for most of the subgroup differences in use—especially the very low rates of cigarette and alcohol use by Blacks. Other possibilities which we have begun to explore, and will continue to examine, are that these relatively low usage rates are strongly influenced by particular religious doctrines and affiliations; differences in parent, peer, and community norms; different attitudes regarding the use of drugs; or differing levels of perceived risk.

The very low rates of drug use (particularly cigarette use and alcohol use) by Black students may be surprising, in light of evidence that drug-related mortality and morbidity are higher among Black than White adults.[1] These contrasting findings are, however, consistent with data from national household surveys comparing Blacks and Whites at various age levels: Black youth show much lower usage rates than White youth; in early adulthood differences are smaller; and by middle adulthood, the drug use/abuse rates are often higher among Blacks.[7,27,28]

It has been observed that "two worlds" of drug use (particularly alcohol use) exist within the Black community: the extremes of abstinence at one end and heavy use/abuse at the other.[27,29] As a result of heavy drug use by a minority (often having limited financial resources, health care, and insurance), Black adults are disproportionately represented in morbidity and mortality statistics, and data from public treatment centers. Such data, rather than findings from general population surveys, often serve as the basis for

Appendix

Measures

The use of alcohol and the use of each of the illicit drugs are measured by questions having the following format: On how many occasions (if any) have you used [name of drug category] . . .

a) . . . in your lifetime?

b) . . . during the last 12 months?

c) . . . during the last 30 days?

Seven response categories are available: 0 occasions; 1–2; 3–5; 6–9; 10–19; 20–39; 40 or more occasions. Annual and monthly prevalence refer to any use of the specific substance during the last 12 months, or the last 30 days, respectively. Daily prevalence refers to use on 20 or more occasions in the last 30 days. An additional question about heavy use of alcohol asks respondents how many times in the last two weeks they had five or more drinks in a row. Cigarette use is measured by a question which asks about use in the past 30 days (none, less than 1 cigarette per day, 1–5 cigarettes per day, about ½ pack per day, about 1 pack per day, about 1½ packs per day, 2 or more packs per day). There is a good deal of inferential data in support of the validity of such self-report measures of drug use, summarized elsewhere.[2,21-25]

The measure of racial/ethnic identification is based on the questionnaire item: "How do you describe yourself? 1. American Indian; 2. Black or Afro-American; 3. Mexican American or Chicano; 4. Puerto Rican or other Latin American; 5. Oriental or Asian American; 6. White or Caucasian; 7. Other."

conclusions about drug use among Black Americans. Our present findings clearly suggest that such data bases are too limited, and can yield a false picture of the subgroup as a whole.

But it is equally true that general population surveys do not provide an adequate basis for generalizations about the full range of drug use problems among Blacks, Whites, or any other racial/ethnic subgroup. Persons in drug treatment, the homeless, and those in prison are not represented in general population surveys. Additionally, dropouts, youth who are chronically absent, those in juvenile detention centers, and those who are homeless, are not well represented in samples of high school students. Here again, because minority group members make up disproportionate shares of most or all of these "high risk" populations, and because the levels of drug use in these populations are greater than average, it is widely believed that drug use is more pervasive among minorities than among Whites. While that remains a possibility among the dropouts, it now seems quite unlikely among young people who complete high school.

In sum, we are confronted with (at least) two worlds of drug use *data*. On the one hand, the findings from general population and school-based surveys clearly and consistently show relatively low levels of drug use by most non-White youth, especially Black Americans and Asian Americans. On the other hand, the public health statistics on mortality, morbidity, and treatment provide a somewhat different perspective. So perhaps our most important conclusion must be that neither form of data provides a complete picture of drug use. However, that should not overshadow the other important conclusion: the majority of non-White youth do complete high school, and among most of these individuals, usage rates for both illicit and licit drugs are lower than average. ☐

Acknowledgments

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Public Health Briefs

Elevated Nicotine Levels in Cervical Lavages from Passive Smokers

Carol J. Jones, PhD, Mark H. Schiffman, MD, MPH, Robert Kurman, MD, Peyton Jacob III, PhD, and Neal L. Benowitz, MD

ABSTRACT

One hundred forty-five nonsmokers found to have normal cytologic diagnoses on routine Pap smears were interviewed regarding environmental exposure to tobacco smoke, and a 3 ml saline lavage of the cervix was collected for measurement of cervical nicotine levels by gas chromatography-mass spectroscopy. Nicotine levels tended to be highest among women exposed to tobacco smoke in the home, intermediate in women exposed only outside the home, and lowest in women recalling no exposure ($p = 0.001$). (*Am J Public Health* 1991;81:378–379)

Introduction

Women exposed to environmental tobacco smoke show biochemical evidence of smoke absorption in their saliva, serum, and urine.[1] A study of four women[2] recently reported that the cervical mucus of passive smokers might contain nicotine and its major metabolite, cotinine, concordant with the finding of increased concentrations of nicotine and cotinine in the cervical mucus of active smokers.[3,4] To confirm this report, we measured the concentrations of nicotine in cervical lavages taken from 145 cytologically normal, nonsmoking women, comparing concentrations in women passively exposed to tobacco smoke with those not exposed.

Methods

Subjects were part of a larger study of cervical neoplasia conducted in three Washington, DC area hospitals.[5] One hundred forty-five nonsmokers found to have normal cervical cytologic diagnoses on routine Pap smears were interviewed regarding environmental exposure to tobacco smoke, and a 3 ml saline lavage of the cervix was collected as previously described.[4] The interview included questions on the following topics: exposure to environmental tobacco smoke inside and outside the home in the previous 24 hours, time since last exposure, number of active smokers in the home, the smokers' relationship to the subject, and their usual intensity of smoking and products smoked (cigarette, cigar, pipe). The questionnaire focused on the 24 hours preceding specimen collection because previous research had indicated that cervical nicotine levels were correlated with active smoking during this period.[4] Specimens were kept frozen at −70° C until tested. All laboratory analyses were performed without knowledge of the smoking histories.

Nicotine was extracted from lavage samples by a modification of a published procedure.[6] The extracts were analyzed by combined gas chromatography-mass spectroscopy,[7] using a stable isotope-labeled analog, nicotine-d_4, as an internal standard. As a quality control measure, 17 aliquots of a pool composed of 19 cervicovaginal saline lavages, 17 from nonsmokers and two from smokers, were included among the specimens to be analyzed, with labels indistinguishable from the test specimens. The quality control data suggested that nicotine was measured reliably in the lavages. The mean of the 17 replicate measurements was 2.72 ng/ml with a coefficient of variation of 13.6 percent and a range of 2.2–3.5 ng/ml.

To classify subjects with regard to environmental tobacco smoke exposure, the following categories were formed: group 1 included women who reported exposure inside their homes within the past 24 hours, group 2 was comprised of women reporting exposure only outside their homes within the past 24 hours, and group 3 included women who recalled no exposure to environmental tobacco smoke in the previous 24 hours.

Address reprint requests to Mark H. Schiffman, MD, MPH, Environmental Epidemiology Branch, National Cancer Institute, Executive Plaza North 443, Bethesda, MD 20892. Dr. Jones is also with EEB/NCI; Dr. Kurman is with Georgetown University Hospital; Drs. Jacob and Benowitz are with the University of California at San Francisco. This paper, submitted to the Journal December 15, 1989, was revised and accepted for publication September 5, 1990.

TABLE 1—Nicotine Levels in Cervical Lavages from 145 Nonsmokers, according to Exposure to Environmental Tobacco Smoke over the past 24 Hours

Nicotine level (ng/ml)	Group 1 (exposed at home)		Group 2 (exposed only outside home)		Group 3* (not exposed)	
	#	%	#	%	#	%
<0.2	5	15.6	6	14.3	22	31.4
0.2–0.3	7	21.9	14	33.3	21	30.0
0.4–0.7	4	12.5	10	23.8	20	28.6
>0.7	16	50.0	12	28.6	7	10.0
Total	32	100.0	42	100.0	70	100.0

Chi-square = 22.5, 6 degrees of freedom, p = 0.001

*Group 3 excludes one woman with an inadequate specimen.

Results

The study population was predominantly Black (63.5 percent); the median age was 30 years (range 16–66). Neither age nor race was associated with environmental tobacco smoke exposure.

A total of 32 women reported exposure to environmental tobacco smoke in the home (group 1), 42 reported exposure only outside the home (group 2), and 71 were unexposed (group 3). In group 1, 24 reported one active smoker in the home, five reported two active smokers, and three reported three or more active smokers in the home. Seven in group 1 recalled additional exposure to tobacco smoke outside the home.

Nicotine values ranged from below the limit of detection (<0.2 ng/ml) to 8.2 ng/ml. Nicotine levels tended to be highest among women in group 1 (exposed in the home), intermediate in group 2 (exposed only outside the home), and lowest in group 3 (not exposed). The median values for the three groups were: group 1, 0.8 ng/ml (range, <0.2–8.2), group 2, 0.4 ng/ml (range <0.2–5.2), and group 3, 0.2 ng/ml (range <0.2–3.8).

For grouped analysis, the nicotine levels were divided into approximate quartiles as shown in Table 1. The association of recalled environmental tobacco smoke exposure in the past 24 hours and detectable nicotine in cervical lavages was highly significant (p = 0.001).

Among the 38 women with exposure at home in the past 24 hours, there was no association observed between nicotine levels and the number of smokers to whom the women had been exposed, the usual smoking intensity of those active smokers, or the number of hours since last exposure. Moreover, for women exposed to smoke in the home, additional exposure outside the home was not related to nicotine level. Cigarette smoke accounted for virtually all of the exposure at home. Among women exposed only outside the home, there was no relationship observed between time since last exposure to environmental tobacco smoke and nicotine level.

Discussion

We observed a significant association between self-reported exposure to environmental tobacco smoke within the previous 24 hours and nicotine levels in cervical lavages from a group of 145 nonsmoking women. Exposure in the home resulted in the highest nicotine levels observed; however, women exposed only outside the home also had elevated levels of nicotine compared to those not exposed.

The biological significance of cervical nicotine levels resulting from passive exposure to tobacco smoke is not known. The levels of nicotine observed in nonsmokers most heavily exposed in our population were on average much lower than those seen in cervical lavages taken from active smokers. We found the median nicotine level in nonsmoking women exposed to environmental tobacco smoke at home to be 0.8 ng/ml (mean 1.1), ranging up to 8.2 ng/ml. We were unable to relate nicotine levels to more subtle distinctions, such as the intensity of home exposure. In contrast, using similar specimen collection and analytic methods, we observed a median value in 31 active smokers of 11.8 ng/ml (mean 34.3, range 2.8–383.4, unpublished data) and were able to correlate increasing intensity of active smoking with increasing cervical nicotine levels (Spearman's r = 0.46, p = 0.009). It is difficult to reconcile the relatively low levels of nicotine that we observed in passive smokers with the large magnitude of the relative risks reported by Slattery, *et al*, who found a risk from passive smoking similar to that of actively smoking women.[8]

It is remarkable, nonetheless, that even a relatively crude exposure categorization provided by questionnaire responses revealed a clear, albeit small, elevation of nicotine levels in the cervical fluid of passive smokers. This supports the idea that even low-level exposure to tobacco smoke in the environment may result in systemic effects, and suggests that passive smoking should be evaluated further as a risk factor for cervical disease. ☐

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Ascertainment of Maternal Deaths in New York City

Machelle H. Allen, MD, Wendy Chavkin, MD, MPH, and Joani Marinoff, PA

ABSTRACT

Maternal deaths in New York City are defined as deaths from any cause in a woman while pregnant or within six months of pregnancy termination. Pilot studies seeking to improve maternal death ascertainment found that selected medical examiner reports contributed an additional 10.5 percent of the total maternal deaths, vital statistics review contributed 6.3 percent, linkage of death tapes of women of reproductive age to live birth and fetal death tapes contributed 1.0 percent. Medical examiner cases should be incorporated into surveillance data for accurate ascertainment of pregnancy associated deaths. (*Am J Public Health* 1991;81:380–382)

Introduction

Maternal mortality, like infant mortality, serves as a sensitive indicator of the level of health and health care in a community. Speckhard, *et al*,[1] reviewed death certificates and medical records in Puerto Rico and found that the maternal mortality rate was increased by a factor of four in 1979. Starzyk, *et al*,[2] increased the ascertainment of maternal deaths in Washington state by 68 percent through linkage of birth and fetal death files with death certificates for all reproductive-aged women. In Georgia, an additional 27 percent of maternal deaths were identified via a similar vital record linkage system.[3] A conservative pooled estimate of the degree to which maternal deaths are underreported in the United States is by a factor of one-third.

In an effort to verify the efficacy of New York City's present ascertainment methods, which relies upon manual review of all death certificates and clinical records, we undertook three pilot studies to augment routine surveillance.

Methods

Routine Surveillance in New York City

New York City is one of 14 areas in the United States that includes a specific question, item #29 on the death certificate, to identify a pregnancy-associated death (Figure 1). The front of the New York City death certificate also lists coded items: ICD Code, Operative Code, Conditions of Interest Code. If pregnancy is mentioned anywhere on the death certificate, it should be reflected in one or more of the above codes. Routine surveillance relies upon manual review of all death certificates. If an associated pregnancy is indicated on the death certificate, hospital records and autopsy reports are requested and reviewed by an obstetrician. In New York City a maternal death is defined as the death of any woman dying of any cause while pregnant or within six months of termination of a pregnancy.

Death Certificate Review

All death certificates of women aged 10–49 who had died 1983–84 were manually reviewed in order to identify those deaths whose immediate cause was *not* coded in the ICD 630-676 range (complication of pregnancy, delivery or puerperium) but for which the coded items indicated association with pregnancy. All previously unidentified maternal deaths were verified by reviewing the hospital charts and autopsy reports. Additionally, the death certificates of previously reported maternal death cases were reviewed to assess whether all four items related to pregnancy had been completed. This review focused solely on the operative code and the conditions of interest code. It did not investigate those deaths where the ICD code might indicate a medical complication of pregnancy. This issue was addressed by reviewing selected medical examiners reports as will be described subsequently.

Linkage

The New York State Department of Health linked fetal death files and live birth files from 1983 and 1984 with death files of all New York City resident females ages 10–49 for these same years. Linkage was based on mother's last name, infant's last name or mother's first name being the same on both maternal and infant documents. All potential matches were verified for decedent's maternal status by the authors' review of available hospital charts and autopsy reports.

Medical Examiner Report Review

In New York City, there are approximately 3,000 autopsies performed annu-

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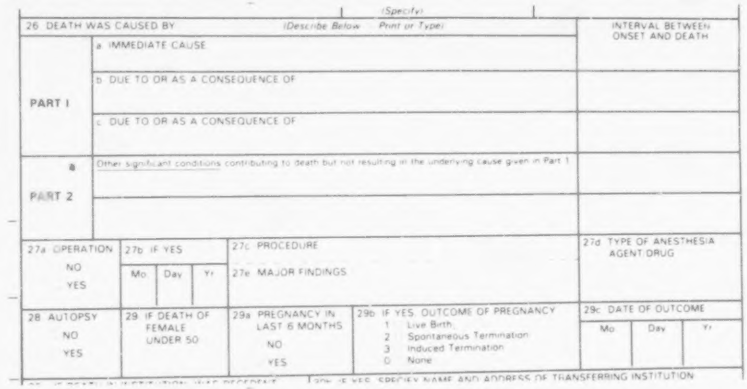
Address reprint requests to Machelle H. Allen, MD, Instructor, Department of Obstetrics and Gynecology, New York University School of Medicine, 550 First Avenue, New York, NY 10016. Dr. Chavkin is Assistant Professor, Columbia University School of Public Health; Ms. Marinoff is Physician Assistant with Planned Parenthood of San Mateo County. This paper, submitted to the Journal January 16, 1990, was revised and accepted for publication September 5, 1990.



(Specify)

26 DEATH WAS CAUSED BY (Describe Below Print or Type)

INTERVAL BETWEEN ONSET AND DEATH

PART I

a IMMEDIATE CAUSE

b DUE TO OR AS A CONSEQUENCE OF

c DUE TO OR AS A CONSEQUENCE OF

PART 2

Other significant conditions contributing to death but not resulting in the underlying cause given in Part 1

27a OPERATION NO YES

27b IF YES Mo Day Yr

27c PROCEDURE

27e MAJOR FINDINGS

27d TYPE OF ANESTHESIA AGENT DRUG

28 AUTOPSY NO YES

29 IF DEATH OF FEMALE UNDER 50

29a PREGNANCY IN LAST 6 MONTHS NO YES

29b IF YES OUTCOME OF PREGNANCY
1 Live Birth
2 Spontaneous Termination
3 Induced Termination
0 None

29c DATE OF OUTCOME Mo Day Yr

SPECIFY NAME AND ADDRESS OF TRANSFERRING INSTITUTION

FIGURE 1—New York City Death Certificate, Item #29.

ally on reproductive aged women by the Medical Examiner. Selected autopsies from 1983–84 were reviewed. A cause of death considered on clinical grounds to have a high probability of being a medical complication of pregnancy (Table 1) as well as drug-related deaths, homicides, and suicides were selected for review. Available hospital charts were subsequently reviewed to confirm whether or not there had been a recent pregnancy.

Results

A summary of maternal deaths identified by the three pilot methods is presented in Table 2. Six additional maternal deaths were found with only one of the coded items as the indicator of a pregnancy (i.e. solely the ICD, operative or condition of interest code reflected a concurrent pregnancy). Review of verified maternal deaths disclosed that Item #29 had not been completed in 40 percent of these cases. When death tapes of women of reproductive age range were matched with live birth and fetal death tapes, five possible matches were identified. However, review of the hospital charts and autopsy reports revealed only one of these women to have had a recent pregnancy. The results of reviewing the medical examiner's reports are presented in Table 3. These reviews yielded 10 additional pregnancy-related deaths, without any overlap with the other methods.

Adjustment of the maternal mortality rates* to reflect these findings involved a 16.6 percent rate increase in 1983, from 40.8 to 51.6/100,000 live births, and a 33 percent rate increase in 1984, from 24.1 to 32.6/100,000 live births. (A retrospective review of maternal mortality rates for New York City through 1987 reveals that the low rate in 1984 represents an unexplained variance, not a trend.)

TABLE 1—Medical Diagnoses that on Clinical Grounds are likely to be Medical Complications of Pregnancy, New York City 1983–84

		Number of Deaths among Women Aged 10–49 Years	
Cause of Death	ICD Code	1983	1984
Septicemia	038	11	12
Diabetes Mellitus	250	58	61
Immune Disorders	279	42	114
Coagulation Defects	286	4	2
Essential Hypertension	401	3	5
Hypertensive Heart Disease	402	17	30
Cardiomyopathy	425	18	10
Subarachnoid Hemorrhage	430	40	32
Intracerebral Hemorrhage	431	51	53
Phlebitis, Thrombosis	451	13	8
Other Venous Emboli and Thrombosis	453	4	3

TABLE 2—Identification of Pregnancy-associated Deaths, by Source, New York City 1983–84

	1983	1984
Current Surveillance Methods	49	29
Death Certificate Review	3	3
Birth Tape Linkage	1	0
Medical Examiner Reports	5	5
Total # Additional Cases	7	7
Revised No. Maternal Deaths	58	37
Percent Increase of Absolute Number	14.3	13.8

Discussion

Review of medical examiner reports proved to be the most useful approach to increasing the ascertainment of maternal deaths in New York City. Routine surveillance identified 82.1 percent of the 95 maternal deaths in 1983 and 1984; review of medical examiner reports contributed 10.5 percent; death certificate review contributed an additional 6.3 percent, linkage contributed 1.0 percent. Unlike Puerto Rico and Washington, ascertainment in New York City was not enhanced by the linkage system. This may reflect the population's high mobility and the alleged use of false names and addresses. A surveillance system utilizing linkage methods will not pick up those deaths associated with therapeutic abortions and ectopic pregnancies since fetal death certificates are not filed in these situations.

Both obstetrical practice and the risk profile of pregnant women have changed over the years. The additional cases identified by these pilot studies reveal some of the impact substance abuse and violence is having on pregnant women and new mothers. Elsewhere we have described human immunodeficiency virus (HIV) related maternal deaths in New York City not detected by routine surveillance.[4]

The results of these pilot studies suggest implementation of a system whereby the medical examiner, or hospital-based pathologist, notifies the maternal mortality surveillance team whenever evidence of a current or recent pregnancy is identified. Physicians filing death certificates need to be reminded to inquire about preg-

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*This rate is not a true rate but rather a *ratio* of deaths per 100,000 live births, i.e. the numerator describes something different than the denominator.

TABLE 3—Identified pregnancy-associated deaths from medical examiner reports by cause of death, New York City 1983–1984.

	# Cases Reviewed			# Pregnancy-Related		
Category	1983	1984	(Total)	1983	1984	(Total)
Medical Conditions	109	224	(331)	3	1	(4)
Drug-related	223	303	(526)	1	3	(4)
Assault, Suicide	102	36	(138)	1	1	(2)
Total	432	563	(995)	5	5	(10)
% Yield				1.15	.8	(1)

nancy history in all women in the reproductive age range and to complete certificates thoroughly.

As this nation has not met the Surgeon General's 1990 goals of no more than five maternal deaths/100,000 live births for any ethnic or geographic group,[5] it is an urgent matter to design preventive strategies. In order to do so effectively, it is imperative to ascertain all cases so as to accurately delineate the risk factors and causes of pregnancy-associated death. It is important to note that the New York City maternal death definition and surveillance is not universal. If a national surveillance system is to function effectively all registration areas need to utilize the same definitions as well as procedures. ☐

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Oral Public Health Programs for the Elderly: 1989

Michael S. Strayer, DDS, MS

ABSTRACT

Forty-eight of the 50 state dental directors, program managers, or state health officials were interviewed by telephone about the nature of current oral health programs for the elderly and factors that aided or hindered the development of such programs. Oral health programs for the elderly were reported in 30 states (63 percent of respondents). A majority of respondents indicated that a perceived or documented need for oral health programs for the elderly and lobbying by local advocacy groups were instrumental in implementing or planning such programs. (*Am J Public Health* 1991;81:382–384)

Introduction

The "graying of America" is a well-accepted phenomenon, as is its effect on the health care system.[1–5] Preventive oral health services and behaviors offer elderly people potential for enhancing the quality of life through improved chewing ability and reduced pain and infection of dental origin, for example. With a decrease in edentulism rates among succeeding cohorts of elderly[6] and an increase in dental utilization,[7] older adults are likely to seek oral health care in greater numbers than did their parents. The U.S. Surgeon General's Workshop on Health Promotion and Aging noted that additional activities are needed in the areas of oral health education and services for older adults,[8] areas falling within the purview of state and local health agencies.

Out-of-pocket reimbursement is the primary method of payment by older adults for dental care.[9] Of the three principal government health care programs, Medicare provides essentially no dental coverage, and the Veterans Administration limits dental coverage to eligible veterans; although Medicaid provides dental coverage for indigent adults in 43 states, 17 states limit benefits to emergency care, denture services, or both.[10]

At the local level, some agencies have begun to allocate resources to the oral health needs of older adults. Kuthy and Odom reported in 1988 that 16 (of 127) local programs provided clinical services in nursing homes and that oral health education programs were offered by 39 (of

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Address correspondence and reprint requests to Michael S. Strayer, Box 201, Postle Hall, 305 W. 12th Avenue, Columbus, OH 43210. Dr. Strayer is with the Section of Community Dentistry, College of Dentistry, The Ohio State University. This paper, submitted to the Journal January 9, 1990, was revised and accepted for publication June 25, 1990.

TABLE 1—Distribution of Oral Public Health Programs for Elderly Persons by Proportion of State Population over 65 Years of Age

	Proportion of State Population Over 65 Years Old		
Program Type	≤10% (n = 14)	11 to 12% (n = 20)	≥13% (n = 16)
Clinical care	3	0	2
Health education	3	5	6
Combination[a]	4	6	1
Total	10	11	9

[a] Both clinical care and hearing education.

125) agencies.[11] Evans found that 35 percent of local dental programs reported providing health education programs at senior centers.[12]

At the state level, no recent information is available on oral health programs for older adults. Miller, *et al,* reported that 50 percent of state health agencies had statutory authority to provide dental health services.[13] A 1983 study of state health agencies identified 20 nonclinical geriatric dental initiatives.[14]

The objectives of this project were to determine the extent and nature of current and proposed oral health programs for older adults and to identify factors that had facilitated or hindered the development of such programs.

Methods

A structured interview was designed to address the study objectives and then revised based on several reviews for clarity and content by a panel of professionals in dental public health, geriatric dentistry, research design, and program evaluation.

All known state dental directors or dental program managers were contacted, first by letter and later by telephone, to arrange a mutually convenient time for the interview. In states without a dental director or program manager, another health department administrator was contacted. All data provided by respondents were considered confidential and are reported here in aggregate form only.

Results

A total of 48 state dental directors, managers, or administrators completed interviews, yielding a 96 percent response rate. Although the interviews were subject to the limitations of self-report surveillance, interviews revealed that some level of oral health programming for older adults existed in 30 states (63 percent of respondents).

A variety of oral health education or promotion activities were provided by 25 of the 30 states, primarily in long-term care facilities. Clinical services were provided to the elderly by 16 of the 30 states, with 3 states having direct responsibility and 10 states offering services jointly with a local agency. In the remaining states, clinical services were provided through arrangements with local educational institutions and dental associations or through loans of portable equipment to private dentists. Of the 18 states without programs for older adults, 6 planned to implement health education or prevention programs within 3 years; 2 of the 6 also planned a clinical component.

In the majority of states with programs, two factors that assisted in the development of programs were identified: a perceived or documented need for oral health programs, and a local advocacy group that lobbied for programs. Obtaining funds that could be used for older adult programs assisted program development in most of the states planning programs.

Low state priority for providing oral health services to older adults and funding restricted to population groups other than older adults were identified by a majority of respondents as hindering the development of older adult programs.

Independent variables including proportion of the state population over age 65, total number of elderly persons in the state, eligibility of elderly to participate in the state Medicaid program, and size of the state dental budget were examined to determine their utility as predictors of a state program. However, none of the independent variables predicted the presence of a state oral health program for the elderly. Table 1 shows the distribution of oral health programs by states in which the elderly population was the same, higher, or lower than the national average of 11 to 12 percent. No evident relation of program and population can be seen.

Discussion

Although changes in dental health care priorities seem inevitable, they must be preceded by careful needs assessment, so that resources are allocated to the most compelling problem areas. To understand the scope of such problems, the needs assessment process should include development of a data base on the dental needs of older adults, both perceived and observed; identification of dental services available to or needed by older adults; and identification of the ranking of dental care on a list of health priorities.

Oral health programs for older adults that were reported to be more comprehensive shared the following three characteristics: a commitment by the state agency to provide oral health programs for older adults, active interest by organized dentistry, and some form of linkage with the local aging network (such as local Area Agencies on Aging or a state Department of Aging). The scope and magnitude of the physical, psychological, and socioeconomic problems encountered by the elderly in obtaining and paying for oral health care make the support of this triad vital to the successful development and implementation of oral health programs for older adults.

A majority of respondents perceived that lobbying by local advocacy groups had assisted in implementing oral health programs and that the lack of such an organized group was a hindrance. This area is one where appropriate assistance must be offered; access to dental consultants with gerontologic training could be extremely helpful to state agencies in their work with diverse advocacy groups. Dental public health can have a significant role in improving the quality of life for the elderly by using the health promotion and prevention skills that improved the oral health of younger populations. ☐

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Effect of Pregnancy during TMI Crisis on Mothers' Mental Health and Their Child's Development

Peter S. Houts, PhD, George K. Tokuhata, DrPH, PhD, Jane Bratz, BS, Mary J. Bartholomew, MS, and Kirby W. Sheffer, MS

ABSTRACT

Five years after the Three Mile Island nuclear accident, the mental health of women who had been pregnant and living within 10 miles of Three Mile Island at the time of the accident was similar to that of women from the same area who became pregnant after the accident. Ratings of the development of the two groups of children when they were 5 years old were also similar. However, women who were pregnant during the crisis and had been "extremely disturbed" about their pregnancies rated their children's health as poorer than did the women who were pregnant later. (*Am J Public Health* 1991; 81:384–386)

Introduction

A review of published research on the psychological and social effects of the Three Mile Island (TMI) accident showed that women living near the facility who were pregnant or had young children were among those who experienced the greatest psychological distress following the accident.[1] A study conducted by the Pennsylvania Department of Health showed little biological effect of the accident on women who were pregnant during the crisis and on their children.[2]

Five years after the accident, the Pennsylvania Department of Health conducted follow-up studies to assess long-term physical and psychological impacts of the TMI crisis on these same women and children. We report here interviews with the mothers concerning their perceptions of their children's health and development and their own mental health.

Method

The study and control cohorts consisted of 7,433 women who gave birth within 2 years of the TMI accident and lived in communities within 10 miles of the nuclear facility at the time of the births. A 50 percent random sample of each of these cohorts was selected: 1,880 mothers who gave birth within 12 months of the TMI accident; and 1,850 mothers who gave birth from 13 to 24 months after the accident. The latter group served as controls. Mothers who had moved out of the area since their children's births, as well as mothers who still lived in the vicinity of TMI, were interviewed. A 97 percent response rate was achieved for each sample,

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Address reprint requests and correspondence to Peter S. Houts, Department of Behavioral Science, Pennsylvania State University College of Medicine, Box 850, Hershey, PA 17033.

Dr. Houts is with the Department of Behavioral Science and Ms. Bartholomew is with the Division of Biostatistics and Epidemiology at Pennsylvania State University College of Medicine. Dr. Tokuhata, Ms. Bratz, and Mr. Sheffer are with the Division of Epidemiology Research of the Pennsylvania Department of Health. This paper, submitted to the Journal October 17, 1989, was revised and accepted for publication May 31, 1990.

TABLE 1—Multiple Regression Analyses Comparing Ratings of Mothers' and Children's Health by Mothers in the Study and Control Groups (N = 2,585)

Dependent Variables	Regression Coefficients for Study-Control Comparisons[a]	Standard Error
Ratings of child's health (high score = poor health)	.0438	.0185
Number of child's visits to physician in past year	.0536	.0317
Developmental profile (high score = poor development)	.0310	.0685
Preschool behavior scale (high score = emotional problems)	.0241	.0211
Ratings of mother's health (high score = poor health)	−.0000	.0174
Mother's depression score	.0151	.0095
Mother's anxiety score	.0147	.0088

Notes: Covariants are listed in text. N varies slightly with number of missing values for each dependent variable. [a]Control group = 0, study group = 1.

TABLE 2—Analyses of Covariance of Mothers' Mean Ratings of Their Children's Overall Health by Timing of Pregnancy and Disturbance about Pregnancy

Disturbance about Pregnancy	Pregnant During Crisis	Pregnant After Crisis	P value of Difference
Extremely disturbed	2.17 (N = 327)	1.95 (N = 87)	.0236
Somewhat disturbed	1.88 (N = 667)	1.94 (N = 478)	.3358
Not disturbed	1.69 (N = 388)	1.71 (N = 638)	.7284

Note: Health was rated 1, excellent; 2, very good; 3, good; 4, fair; or 5, poor. Covariants are listed in text.

yielding 1,833 pairs of mothers and children in the study sample and 1,808 pairs in the control sample.

The analyses reported here involved women who were pregnant at the time of the accident, i.e., those who gave birth during the 9 months after the accident (study group), and those who gave birth during the corresponding 9 months a year later (control group). The study group was interviewed in 1985 when the index children were 5 years old, and the members of control group were interviewed in 1986 when their children were also 5 years old.

The dependent variables analyzed were the number of physician visits,[3] the developmental profile for 5-year-old children,[4] the preschool behavior scores suggestive of emotional problems,[5,6] the mothers' anxiety and depression scores,[7,8] and the mothers' overall ratings of their children's and their own health.

Results

We fitted multiple regression models with terms for study versus control group, month of child's birth, child's sex, child's birth order, mother's marital status, time mother spends caring for child, mother's education, mother's employment status, mother's race, mother's age, household size, and distance of home from TMI at the time of the crisis. Dependent variables were log transformed to adjust for skewness.

Table 1 shows that all but one coefficient (mother's health) are in the direction suggesting poorer status for the study group. Only one coefficient, however, mothers' overall ratings of their children's health was statistically significant. Women who gave birth within 9 months of the accident gave their children's health poorer mean ratings than women who gave birth in the same months a year later.

Examination of the distribution of these general health ratings showed that the difference in the two groups was largely in the percentage of mothers who rated their children's health as "excellent"; that percentage was 7 percent lower among the study mothers than among controls. We also examined interactions among the study and control groups' ratings and mothers' age, birth order of the child, and mothers' education while controlling for the same 11 covariants. No interactions were statistically significant.

Mothers were grouped based on how disturbed they said they had been about their pregnancies when they were interviewed 5 years earlier, shortly after their children's births (Table 2). The poorer health ratings were primarily among mothers who had been "extremely disturbed" about their pregnancies. The statistic for this interaction effect had a probability of .0635. The difference between the two groups was significant only for the "extremely disturbed" group.

Discussion

The interview data reported here are primarily subjective and indicate how the experience of being pregnant during the TMI crisis affected the mothers' views of their children and themselves. Although the findings showed a poorer status for women who were pregnant during the crisis, the differences between this group and the controls were small. The only notable difference was in the smaller proportion of study group mothers who rated their children's health as "excellent." Small but statistically insignificant differences between women who were pregnant during the crisis and women who became pregnant afterward were found for long-term anxiety, long-term depression, ratings of the children's development, and reports of children's physician visits.

Women who were pregnant during the crisis and who were also "extremely disturbed" about their pregnancies tended to rate their children as less healthy. This pattern is similar to that reported by Roht, *et al*, who found that people living near toxic waste sites were more likely to report health problems if they believed that the sites could affect their health.[9] Roht, *et al*, interpreted their findings as an example of response bias induced by the expectations of the respondents.

The present findings can be interpreted similarly: that is, the mothers of children born within 9 months of the TMI accident, who were "extremely disturbed" because of possible long-term effects of radiation released during the accident, may have paid special attention to their children's health problems, resulting in the general impression that their children were less healthy. Whether these children were less healthy than controls will be investigated further in a medical record review study now in progress. ☐

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Using a State Cancer Registry to Increase Screening Behaviors of Sisters and Daughters of Breast Cancer Patients

Peter S. Houts, PhD, Sandra L. Wojtkowiak, BSN, Mary A. Simmonds, MD, Gene B. Weinberg, DrPH, and Daniel F. Heitjan, PhD

ABSTRACT

The Pennsylvania Cancer Registry was used to contact breast cancer patients and, through them, their adult sisters and daughters. The sisters and daughters were counseled concerning their higher than average risks for breast cancer and their need for mammography and breast self-examination. Results showed a 9 percent increase in mammography and a 10 percent increase in breast self-examination rates for the counseled over control group. Costs were $49 per counseled sister or daughter indicating a need to increase cost effectiveness before implementation is practical. (*Am J Public Health* 1990; 81:386–388)

Introduction

Cancer registries have contributed to the treatment and prevention of cancer by making it possible to study the distribution of the disease, its causes, and the planning of services.[1,2] Women whose mothers or sisters have had breast cancer have two to three times the usual risk for developing breast cancer in their lifetimes.[3] We used the registry to contact breast cancer patients and through them first degree female relatives in order to counsel them about what to do to ensure early diagnosis and treatment of breast cancer, should it be found.

Methods

Potential participants were sisters and daughters of breast cancer patients diagnosed between June 1982 and June 1984 and living in eight counties in central Pennsylvania. Names of all breast cancer patients from four counties were obtained from the Pennsylvania Cancer Registry at the beginning of the study and randomly assigned to three groups where sisters and daughters were to be counseled in years 1, 2 and 3, respectively, of the project. At the end of the first year, four additional counties were added to the study and randomly assigned to the years 2 and 3 groups. At the end of 21 months, the knowledge, attitudes and behavior of sisters and daughters of breast cancer patients in all three groups were assessed using telephone interviews. Those who had received counseling in the first 21 months were "experimental," while those yet to be counseled were controls. This design controlled for the effects of media attention and public education programs that occurred during the first 21 months of the study.

Permission to contact the patients was obtained from the physicians listed on the cancer registry forms. When permis-

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From the Department of Behavioral Science (Houts), the Department of Medicine (Simmonds), and the Center for Biostatistics and Epidemiology (Heitjan), Pennsylvania State University College of Medicine; Community Health Services of Hanover and Spring Grove (Wojtkowiak); and the Pennsylvania Department of Health (Weinberg). Address reprint requests to Peter S. Houts, PhD, Department of Behavioral Science, Pennsylvania State University College of Medicine, Box 850, Hershey, PA 17033.

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TABLE 1—Characteristics of Study Populations

Patients	Experimental	Control
Total number selected[a]	611	395
MD permission given	431 (71%)	221 (56%)
Pts who agreed to participate	302 (70%)	165 (75%)
Number with eligible sisters or daughters[b]	160	111
Mean patient age[b]	62	61
Stage of disease at diagnosis[b]		
Stage I	4%	4%
Stage II	70%	53%
Stage III	25%	40%
Stage IV	1%	3%
Year of diagnosis[b]		
1982	27%	14%
1983	56%	48%
1984	17%	38%
Sisters and Daughters		
Total number participating	370	242
daughters	46%	48%
sisters	54%	52%
Mean age	48	46
Education		
Less than high school	26%	24%
High school graduate	45%	48%
Post high school	29%	28%

a) Excludes patients who died before evaluation interviews or were living outside of Pennsylvania.

b) Excludes patients without sisters or daughters, or where sisters and daughters were less than 18 years of age or were living outside Pennsylvania.

TABLE 2—Sisters' and Daughters' Knowledge and Behaviors Concerning Breast Cancer[a]

	Experimental Group	Control Group	Estimated Treatment Effect (95% CI)[e]
Information[b]			
Mean percent of correct answers on general information about breast cancer	71%	65%	6.83% (4.51, 9.15)
Mean percent of correct answers on signs of breast cancer	30%	24%	6.75% (4.61, 8.90)
Breast Self-examination[b]			
Confidence in ability to do BSE (0 = not at all, 3 = very)	2.23	2.04	.20 (.07, .32)
			Relative Odds (95% CI)[f]
Percent doing BSE monthly or bimonthly	49%	39%	1.53 (1.08, 2.17)
Mammography[c]			
Percent ever having a mammogram	57%	52%	1.36 (0.91, 2.03)
Percent who had a mammogram since receiving education[d]	39%	30%	1.57 (1.03, 2.40)

a) Controlling for the index patient's stage at diagnosis and year of diagnosis and the respondent's age, education and whether she was a sister or daughter.

b) Experimental group N = 369, Control group N = 242.

c) Only sisters and daughters ages 35 and older. Experimental group N = 269, control group N = 185.

d) Experimental group women were asked if they had a mammogram since their telephone education and the control group was asked the date of their most recent mammogram. Persons in the control group were scored positively if they had a mammogram in the past 21 months, which was when experimental group education occurred.

e) Continuous outcome: additional effect of treatment on mean response when all other factors are held constant.

f) Dichotomous outcome: relative odds for treated versus control respondents when all other factors are held constant.

sion was given, patients were contacted and asked for access to their sisters and adult daughters living in Pennsylvania who were then reached and counseled by telephone. Counseling consisted of asking questions about their knowledge, attitudes, and behaviors related to breast cancer and, where responses were not consistent with American Cancer Society (ACS) guidelines, explaining and discussing the reasons for the guidelines. The counselors urged the sisters and daughters to follow ACS guidelines because of their risks for future breast cancer and the excellent prognosis when cancer is discovered at an early stage.

All sisters and daughters were invited to attend education sessions held throughout the eight-county study area and were encouraged to come in family groups. The education sessions included explanations of breast cancer risks, the importance of breast self-examinations (BSE) and mammography as well as opportunities to practice BSE on model breasts and on themselves.

Results

Characteristics of the experimental and control groups are shown in Table 1. The control group had a lower physician permission rate, fewer early stage diagnoses, and more recent diagnoses. One-third of experimental group women attended the voluntary teaching sessions in their communities.

Table 2 reports estimated effects of the experimental treatment on general knowledge of breast cancer, breast self-examination, and mammography. Multiple linear regression was used to make comparisons between the groups based on continuous outcomes, and multiple logistic regression was used to make comparisons based on dichotomous outcomes. All analyses adjust for index patient's stage at diagnosis and year of diagnosis and the respondent's age, education, and whether she was a sister or a daughter. Treatment effects were parameterized as the additional effect of treatment on the mean response for continuous outcomes, and as the relative odds for treated versus control for dichotomous outcomes. As the table shows, all comparisons favored the treatment group, although the effects were not large.

We also investigated a number of possible interactions including: whether those who attended the group sessions showed a greater change in screening behaviors than those who did not attend; whether sisters responded differently to the program than daughters; whether higher socioeconomic status women responded differently than lower socioeconomic status women; whether younger women responded differently than older women; and whether the index patient's stage of disease at diagnosis was related to the program's impact. None of these variables interacted significantly with the experimental/control variable in their effects on the sisters' and daughters' breast self-examination, ever having a mammogram, or having a mammogram since receiving education. We also investigated whether the eight counties were significantly different on any of the outcome measures. They were not.

Cost estimates, based on hours spent by project personnel, were $49 per sister or daughter counseled. Twenty-five percent of these costs were expended in obtaining physician permissions, 50 percent in obtaining patient permissions and counseling relatives, and 25 percent was for supervision and secretarial support.

Discussion

The medical and demographic differences between the experimental and control groups may be the result of when subjects were recruited. Since the control group was recruited last, they had the greatest gap between diagnosis and recruitment which may have contributed to their lower physician permission rate and lower participation among patients with early stage at diagnosis.

The findings indicate that the program did affect participants' knowledge, attitudes and behaviors, although the impact on behavior was generally modest. Moreover, only half of the eligible sisters and daughters were counseled due to high physician (29 percent) and ~~patient (30~~ percent) refusal rates. At the time of the evaluation interviews, 43 percent of counseled sisters and daughters over age 35 had never had a mammogram. Similar findings have been reported with other groups at high risk for cancer.[4,5]

The control group mammography rate of 52 percent is also of interest because it is about 10 percent higher than has been reported for the population as a whole[6,*] suggesting that a substantial number of these high-risk women had already taken action to deal with their increased risks. This may have limited the potential impact of the program.

If this program were implemented statewide in Pennsylvania, assuming the same physician and patient refusal rates, the cost would be approximately $227,000 yearly. If, as in this study, there was a 9 percent increase in mammogram rates and, assuming a 1 percent yield of cancers detected (double the usual rate because of working with a high risk population), four cancers would be discovered earlier than otherwise at a cost of $56,963 per diagnosis. The real costs may be lower than these estimates since the excess in mammography rates could be assumed to occur in subsequent years as well. Furthermore, costs of recruiting physicians might also be lower in future years because physicians' increased familiarity and experience with the program might result in more responses to first requests, thereby requiring less follow-up effort. An alternative approach could be to ask physicians to voluntarily refer high-risk patients to the education program. This would reduce recruitment costs and possibly increase patient participation, but would make the program more dependent on physician cooperation. We conclude that it is feasible to use a state cancer registry to contact and counsel first degree relatives of cancer patients, but that additional work is needed to increase the program's cost effectiveness. ☐

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Cost-Effectiveness of Alpha-1 Antitrypsin Replacement Therapy in Treatment of Congenital Chronic Obstructive Pulmonary Disease

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Joel W. Hay, PhD, and Eugene D. Robin, MD

ABSTRACT

Background: Alpha$_1$-antitrypsin (AAT) replacement therapy is an expensive intervention ($20,000–$30,000 per patient annually) which may slow or arrest the progression of chronic obstructive pulmonary disease (COPD) in AAT-deficient patients. While FDA-approved, therapy efficacy is unknown. The costs and benefits of AAT replacement therapy were evaluated for patients with congenital COPD.

Methods: Epidemiological and disease cost data were taken from published sources. A discrete-time model of disease stage probability transition was developed to calculate the present-value expected cost of disease treatment, under a range of possible therapy efficacy and other parameter values.

Results: At an efficacy of 70 percent, the cost per life year saved with AAT replacement therapy would be between $28,000 and $72,000 (1990 US dollars), depending on patient age, sex, and smoking status. At 30 percent efficacy, the cost per life year saved range would be between $50,000 and $128,000. A controlled efficacy study would cost $53 million or less, if the true efficacy were 50 percent or better.

Conclusions: With efficacy of 30 percent or higher, therapy cost-effectiveness would be comparable to other widely used medical interventions. The economic assessment methodology was used to evaluate both the therapeutic innovation and the value of additional clinical research. (*Am J Public Health* 1991; 81:427–433)

Introduction

More than 10 million Americans are afflicted with chronic obstructive pulmonary disease (COPD). It has been estimated that roughly 1 percent of this condition is caused by congenital alpha$_1$-antitrypsin (AAT) deficiency.[1] AAT is a human protease inhibitor which acts in the interstitial lung tissues to prevent damage caused by neutrophil elastase. In addition to COPD, AAT deficiency leads to other clinically significant disorders, including hepatic complications in about 10 percent of patients.[2,3]

In a prospective Swedish study, 8 percent of 183 AAT-deficient eight-year-old children had asthma or bronchitis in the previous year.[4] The mean age for onset of dyspnea, presumably related to COPD itself, was 40 to 45 years in patients with AAT deficiency. The mean life expectancy was 50 to 65 years. Smokers with AAT deficiency experienced a more rapid clinical course, with age at onset of symptoms and age at death advancing by more than a decade relative to nonsmokers.[3]

Human alpha$_1$-proteinase inhibitor (Prolastin) has been approved by the Food and Drug Administration (FDA) for replacement therapy in emphysema patients with congenital AAT-deficiency. Because of the small potential market, it has been granted "orphan drug" status. AAT replacement therapy has been shown to improve the levels of AAT in serum and epithelial tissue.[5] The clinical efficacy of AAT replacement therapy in slowing congenital COPD progression is not yet known.

Unusual policy issues are raised in examining the efficacy and cost-effectiveness of this therapy, with its small number of potential beneficiaries. Because a prospective randomized clinical trial to determine outcome efficacy could cost over $1 billion (see below), health care providers and patients must use other means to decide whether or not to make use of a medication of high cost and unknown benefit. This lack of detailed data on therapeutic risk and benefit arises often, since many forms of medical therapy are used in the absence of controlled clinical trials.

Despite the uncertainties regarding therapeutic outcomes, new immunoassays for AAT-deficiency blood screening are under development (E. Silverman, personal communication, January 1990). Several public health agencies are examining population screening, including the state of Oregon, which is considering adding AAT-deficiency screening to its list of required blood tests for newborns (S. Buist, personal communication, May 1990).

No prior studies have specifically addressed the medical costs associated with AAT deficiency. Previous studies have found the medical and economic costs of COPD to be substantial,[5,7] Oster et al. reported the average medical costs per patient year to range from $587 to $6,238 (1990 dollars) for COPD patients, depending on the number of years after onset of symptoms.[7] In a prospective study of the medical costs for treatment of COPD, Strauss *et al*[8] and Bergner *et al*[9] reported costs averaging $14,647 per year (1990 dollars). The Bergner *et al* study sample was limited to homebound patients with both a

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Address reprint requests to Joel W. Hay, PhD, Senior Research Fellow, The Hoover Institution, Stanford University, Stanford, CA 94305-6010. Dr. Robin is Professor Emeritus, School of Medicine, Stanford University. This paper, submitted to the Journal February 13, 1989, was revised and accepted for publication August 13, 1990.

Editor's Note: See also related editorial p. 414 this issue.

TABLE 1—Cost Effectiveness of AAT Replacement Therapy for Smokers and Nonsmokers

Age	Life Expectancy (undiscounted years)*	Present Value of Average COPD Medical Costs ($1000s)	Present Value of AAT-Replacement Therapy Costs ($1000s)	Therapeutic Increase in Life Expectancy (undiscounted years)	Cost per Life Year Saved ($1000s)
PiZ Males					
Smoking					
30	13.97	$ 82	$456	22.42	$33
40	9.78	$ 96	$403	18.91	$33
50	6.69	$106	$335	14.99	$32
60	7.14	$ 85	$287	9.28	$39
70	2.59	$109	$182	7.99	$28
Nonsmoking					
30	32.06	$ 27	$504	9.76	$64
40	24.47	$ 34	$459	8.63	$64
50	17.93	$ 39	$398	7.12	$68
60	12.82	$ 38	$328	5.31	$72
70	7.45	$ 48	$236	4.59	$57
PiZ Females					
Smoking					
30	17.79	$ 71	$485	24.09	$33
40	11.96	$ 89	$436	21.45	$32
50	9.28	$ 95	$382	16.88	$34
60	7.68	$ 95	$317	12.05	$36
70	5.50	$ 92	$242	8.33	$36
Nonsmoking					
30	32.43	$ 29	$520	13.84	$50
40	22.69	$ 47	$472	13.94	$45
50	13.79	$ 74	$400	13.73	$41
60	10.66	$ 74	$336	9.97	$44
70	7.27	$ 71	$258	7.09	$43

*Source: Adapted from Larsson[3]

COPD diagnosis and Forced Expiratory Volume (FEV_1) scores below 60 percent of normal, a severely impaired group.

Using a societal perspective to measure costs and benefits, we applied available data on congenital COPD incidence, prevalence, patient survival, and costs of COPD medical treatment to cost-effectiveness evaluation of therapy. By examining the cost-effectiveness of AAT replacement therapy under alternative efficacy assumptions, we consider whether determination of therapeutic efficacy would itself be cost-justified. Possible strategies for population screening of congenital AAT-deficiency are also briefly considered.

Methods

We estimated costs per life-year saved (CLYS) for homozygous (PiZ) AAT-deficient individuals: the ratio of the change in expected net COPD medical cost resulting from AAT replacement therapy relative to the change in years of life saved through the intervention (see Appendix).[10–12] Future costs and benefits were discounted to current age using a discount rate of 5 percent.[11–14]

The best available evidence on the natural history of disease in AAT-deficient individuals was reported by Larsson, based on 1 to 14-year follow-up data for 246 Swedish adults with PiZ AAT-deficiency.[3] Column one of Table 1 adapts Larsson's survival data to the adult US population using 1986 US vital statistics. Early estimates of the median survival subsequent to diagnosis of COPD were consistently around five years among subjects with an initial FEV_1 close to 1.0 liter.[15,16] A more recent study by Burrows *et al* shows longer overall survival, but with substantial differences across COPD subcategories.[17] In atopic subjects or nonsmokers with known asthma, the ten-year mortality rate (not age adjusted) was 15 percent, while for smokers with an emphysematous form of COPD, the ten-year mortality rate was 60 percent. Since the mean age in the Burrows *et al* study population was 63, the ten-year survival for the milder forms of COPD was presumably little different from that for non-COPD individuals. The eight-year median survival for emphysematous smokers reported by Burrows *et al* probably more closely matches the prognosis for AAT-deficient patients after COPD diagnosis. While long-term use of therapy might improve or impair the quality of life, we use change in survival as the prime criterion for therapeutic efficacy and outcome assessment.

COPD Medical Care Costs

Several studies show that AAT-deficient individuals tend to develop relatively severe forms of COPD.[1–3] We used the Bergner *et al* and Strauss *et al* medical care cost estimates.[8,9] In their multivariate estimates, annual costs of care did not change with physician specialty, type of hospital (federal or community), insurance coverage, age, sex, income, entry FEV_1 score, length of study participation, or death of patient.[8,9]

The Bergner *et al* and Strauss *et al* study subject inclusion criteria, patient survival, and patient demographic characteristics matched those of the Burrows *et al* study fairly closely, suggesting that a six- to 10-year median COPD patient survival with annual medical costs of $12,000 to $16,000 (1990 dollars) would characterize the course of COPD after diagnosis in AAT-deficient individuals. Conservatively, since no estimates are reported in the recent COPD cost studies, we ignored diagnostic and treatment costs associated with respiratory disease prior to development of FEV_1 values below 65 percent of normal, or with other symptoms outside the Burrows *et al* emphysematous subject category definition. As shown below, this downward bias on the medical cost estimates would have little impact on the CLYS estimates.

Table 1, column two shows the present value of lifetime expected COPD medical care costs, estimated by attributing average excess medical costs per year over the eight-year median survival period prior to death for the fraction of excess deaths due to COPD in each age-, sex-, and smoking-adjusted five-year interval for PiZ cohorts (see Appendix).

AAT Replacement Therapy Intervention

Human alpha$_1$-proteinase inhibitor (Prolastin) is prepared from human plasma screened for hepatitis B and human immunodeficiency virus (HIV) antibodies and heat-treated for at least 10 hours. Hepatitis B vaccine is recommended for all

TABLE 2—Cost per Year of Life Saved by AAT-Deficiency Replacement Therapy at Patient Age 40 ($1,000s)

Outcome Efficacy	PiZ Males		PiZ Females	
	Smoking	Nonsmoking	Smoking	Nonsmoking
10%	$107	$352	$107	$214
30%	$ 50	$128	$ 49	$ 83
50%	$ 38	$ 83	$ 38	$ 56
70%	$ 33	$ 64	$ 32	$ 45
90%	$ 29	$ 53	$ 28	$ 38

recipients, and hepatitis B immune globulin is recommended for recipients who have not yet developed adequate antibody titers.

Prolastin is infused intravenously once per week at a dosage of 60 mg/kg. The current medication cost is $0.10 per mg, or $20,000 per year for a 64 kg adult.[18] The safety record for Prolastin has been described as excellent to date.[19] We assumed average annual costs would be $30,000 to include pharmacy markup, physician monitoring, intravenous infusion, and treatment of side effects. We also assumed that, once initiated, therapy would be maintained throughout the patient's life.

We found no data relating Prolastin in treatment of COPD to improvement in patient outcome. Wewers *et al*, in a study of the safety and biochemical efficacy of Prolastin, provided 507 infusions to 21 PiZ homozygous AAT-deficient patients.[5] They found fourfold increases in AAT serum and epithelial lining fluid levels and doubling in anti-neutrophil elastase capacities subsequent to infusion, but these changes may not reflect improved patient outcome.

AAT-deficiency also produces liver disease, hemorrhagic disorders, pancreatic fibrosis, and other less frequent symptoms.[2,3] Prolastin has shown efficacy in treatment of panniculitis in PiZ patients.[20] However, no efficacy has been suggested for Prolastin for other non-COPD disorders associated with congenital AAT deficiency.[2]

It is not known whether congenital COPD can be fully prevented, somewhat slowed, or merely delayed, so it would be inappropriate to assign a specific value for the outcome efficacy of AAT replacement therapy in treatment of congenital COPD and/or other clinical disorders caused by congenital AAT deficiency. Instead, we have opted to define a hypothetical "outcome efficacy" in terms of the percentage of potential normal survival (age-, sex-, and smoking-status-adjusted) that could be salvaged through therapeutic intervention, and to show the cost-effectiveness of AAT replacement therapy at different "outcome efficacy" values ranging from 0 to 100 percent. That is to say, for each five-year age interval, we define the "outcome efficacy" as x percent if the AAT-deficient individual's probability of survival during that period has increased by x percent relative to a normal (non-AAT-deficient) individual (see Appendix). We do not attempt to detail the specific clinical pathways whereby these hypothetical survival benefits would be achieved.

Results

Table 1, column three shows the net present value of lifetime expected AAT replacement therapy costs in thousands of dollars assuming an "outcome efficacy" of 70 percent. Lifetime therapy costs can only be estimated contingent on efficacy level, since lower efficacy levels will reduce the patient's survival probabilities and thus lower the lifetime expected cost of therapy. The figures in column three also net out the savings in medical treatment of COPD due to reduced probability of disease. Column four shows the change in life expectancy associated with a 70 percent "outcome efficacy" for each age-, sex-, and smoking-status-adjusted cohort. Column five provides CLYS estimates for each adjusted cohort under the assumptions described above, including 70 percent "outcome efficacy," 5 percent time discount rate, eight-year median survival from time of COPD diagnosis, $30,000 annual cost of AAT replacement therapy, $14,647 annual medical cost for treatment of COPD, and a healthiness weight of 75 percent (of normal) applied to average reported days of excess restricted activity (see Appendix). The CLYS values ranged from $27,624 to $71,945; for smoking males and females, most CLYS values were at or below $40,000, while for nonsmoking males and females they were $41,000 to $72,000.

It is a bit puzzling that the CLYS would be as low as $28,000 when the annual AAT treatment cost is $30,000. However, based on the Larsson data, compared with normals, the relative risk of death for PiZ smokers is extremely high, ranging from 27:1 for 35-year-old PiZ male smokers to 4:1 for 60-year-olds. For the individuals at highest risk, after discounting future costs and benefits, and netting out the avoided COPD medical costs (Table 1, column two), at hypothetical "efficacy" levels of 70 percent, $30,000 of therapy buys about a year of additional life expectancy.

To put results shown in Table 1, column five in perspective, they are within the range of such values reported for a number of generally accepted medical interventions. For example, adjusting to 1990 US dollars, Drummond reported that treatment of hypertension in 40-year-old males would cost $30,200 per year of life saved, estrogen therapy for postmenopausal women would cost $42,600, single-vessel coronary bypass would cost $57,400, and hospital hemodialysis would cost $85,400.[10] Kinosian and Eisenberg reported that antihypercholesterolemia therapy would cost between $20,500 and $135,000 per year of life saved, depending on heart disease risk factors.[11]

Table 2 shows the impact of alternative "outcome efficacy" assumptions on AAT replacement therapy CLYS values, holding all other parameters at the same values as in Table 1, column five, for 40-year-old adults. At 10 percent "outcome efficacy" the CLYS estimates exceeded $100,000 but at 90 percent "outcome efficacy" they dropped to less than $30,000 for PiZ smokers. Even if AAT replacement therapy were able to improve congenital COPD survival probabilities by a relatively low 30–50 percent, the values would still be within the range calculated for other frequently used medical interventions, particularly for PiZ smokers.

Sensitivity Analysis

Table 3 shows the effects of 40 percent changes in each key model parameter. The CLYS results are fairly stable. Of the five key model parameters shown, only changes in the annual cost of AAT replacement therapy induced as much as a one-for-one percentage change in CLYS values. Thus, with the exception of assumed "outcome efficacy" (Table 2) and therapy cost, the results were not highly sensitive to underlying modeling assumptions.

TABLE 3—Sensitivity Analysis of Base-Case Model CLYS Values: Effects of ±40 Percent Change in Each Parameter

	PiZ Males (Age 40)		PiZ Females (Age 40)	
	Smoking	Nonsmoking	Smoking	Nonsmoking
Base-Case Value	$32,572	$63,869	$31,959	$44,823
	% Change in Base-Case CLYS Values			
Time Discount Rate				
Base Case: 0.05%				
+40%	−8.7%	−14.4%	−9.2%	−12.6%
−40%	10.4%	20.1%	11.6%	17.9%
Median Survival from Onset of COPD				
Base Case:8 YEARS				
+40%	−6.7%	−4.3%	−7.1%	−5.7%
−40%	5.9%	3.8%	6.4%	5.2%
Annual Cost of AAT Replacement Therapy				
Base Case: $30,000				
+40%	44.2%	41.7%	43.7%	42.0%
−40%	−44.2%	−41.7%	−43.7%	−42.0%
Annual Cost of COPD Medical Care				
Base Case: $14,647				
+40%	−4.2%	−1.7%	−3.7%	−2.0%
−40%	4.2%	1.7%	3.7%	2.0%
Restricted Activity Day Quality Weight*				
Base Case: 75%				
+33%	2.5%	4.5%	5.1%	6.7%
−40%	−2.9%	−4.9%	−5.5%	−7.0%

*It is assumed that the upper bound on health value of restricted activity days is 100% (= 133% of 75%).

TABLE 4—Economic Assessment of Randomized Clinical Trial Outcomes

Outcome Efficacy	Expected Five-Year Mortality Difference	Minimum Study Arm Sample Size (Power = 80%)	Outcome Efficacy Study Cost ($ Millions)	Cost per Life Year Saved
10%	0.015	10,200	$1,530	$216,572
30%	0.046	1,000	$ 150	$ 85,543
50%	0.077	350	$ 53	$ 59,024
70%	0.108	150	$ 23	$ 47,324
90%	0.139	80	$ 12	$ 40,440

Should an Outcome Efficacy Study Be Done?

We can now consider the costs and benefits of undertaking a clinical trial to establish efficacy of AAT replacement therapy. It has been estimated that demonstration of a statistically significant therapeutic efficacy in slowing the decline of lung function would require a clinical trial involving 300–500 patients over a five-year study period.[21] Generally, the value of a randomized controlled clinical trial depends on the prior probability distribution of successful trial outcomes, the size of treatment effect needed to show clinical benefit, and the relative rates of outcome events (e.g., deaths) in treatment and control patients.[22–25] In the case of the already approved AAT therapy, there are penalties associated with not establishing therapeutic efficacy irrespective of the true value. If the true outcome efficacy is very high, many potential beneficiaries may be missing an important medical intervention. If the efficacy is low, then it may be that too many dollars are already being devoted to an expensive but ineffective therapy.

Table 4 considers the design parameters of a randomized controlled trial to evaluate AAT replacement therapy efficacy. Based on the PiZ survival estimates described above, the average five-year mortality probability for a PiZ cohort age 30–60, with equal numbers of men and women and 30 percent smokers would be 18.4 percent. Greater study efficiency would result if the study were targeted to higher risk patients by age, sex, or smoking status, but the generalizability of the findings would then be suspect. For each hypothetical "outcome efficacy," the second and third columns of Table 4 show, respectively, the expected five-year mortality difference and the minimum study sample size needed per study arm to achieve 80 percent power (20 percent risk of false negative results) with a two-sided alpha significance level of .05.[24,26]

Table 4, column four, shows the cost of the study under the assumption that each patient in the AAT therapy arm would require five years of therapy at $30,000 per year. Additional study monitoring, data collection and analysis costs would be trivial relative to these therapy costs. Given that the annual NIH research budget is around $7 billion, less than $3,400 per fatality, study samples needed to detect "outcome efficacy" levels below 30 percent would be prohibitively expensive, particularly since the therapy might partially benefit only 50,000 to 100,000 individuals.

The economic analysis implies that clinical and statistical considerations are not sufficient design criteria in this case. Table 4, column five shows the average CYLS from AAT therapy at each "outcome efficacy" level, assuming that it would be impossible to restrict access to medication by age, sex, or smoking status if the therapy were found to be significantly effective in extending life expectancy. As shown, the "outcome efficacy" of AAT therapy would have to be 50 percent or better to be cost-effective relative to other medical therapies and thus to be economically viable. Detecting economically justifiable therapeutic benefits would thus require smaller sample sizes than would be implied by focusing exclusively on clinical decision criteria. Alternatively, to achieve greater efficiency, the clinical trial could be targeted by age, sex, smoking status, or other risk factors, and if successful, the therapy usage indications could be similarly restricted.

TABLE 5—Projected Costs of Screening for PiZ Emphysema in Various Subpopulations (All Figures in Millions)

	Relevant Adult Population	Cigarette Smokers	Subjects with Established Emphysema	Relatives of Patients with Emphysema	Relatives of Patients with Alpha-1 Antitrypsin Globulin
Estimated Size of Population	100	30	10	20	0.5
Cost of Screening (Dollars)	10,000	3,000	1,000	2,000	50
Cost of Follow-up (Dollars)	1,000	300	100	200	166
Number of False Positives	1	0.3	0.1	0.2	0.005

Congenital COPD Screening Strategies

The availability of a new therapy also raises the issue of the cost-effectiveness of screening to detect early disease and to institute therapy to prevent or retard progression. With rare diseases such as AAT deficiency, the assumed screening rate of false positives has a major impact on the value of screening programs. There are two types of false positive measurements. Technical false positives relate to problems in the performance and interpretation of clinical tests. Biological false positives occur in subjects with a given biochemical or genetic abnormality that is never expressed in frank disease. We assumed that the rate of the former was 1 percent, which is conservative for most tests. We also conservatively chose to ignore biological false positives. However, in a small sample, 25 percent of PiZ homozygotes showed little evidence of COPD.[3] In another study, nonsmoking PiZ homozygotes demonstrated substantial variability in FEV measurements and in COPD prevalence, suggesting a high rate of biological false positives.[1]

We examined some alternative populations that might be screened, estimated false positive rates, and evaluated the screening costs (Table 5). We conservatively ignored the follow-up and treatment costs for false positives. Since only 50,000 to 100,000 individuals in the United States are potentially treatable, some of these screening strategies would add significantly to the cost of intervention, even if Prolastin were highly efficacious. Choice of screening strategies should therefore await more precise efficacy calculations.

Discussion

COPD in congenitally AAT-deficient individuals is a severe, life-threatening illness. AAT replacement therapy leads to promising biochemical improvements in serum and lung-fluid measurements. Our analysis showed that if AAT replacement therapy were able to narrow the age- and sex-adjusted survival probability gap between PiZ and normal individuals by at least 30–50 percent, the cost-effectiveness of AAT replacement therapy would be comparable to widely used medical interventions, including antihypertensive medications, coronary bypass, antihypercholesterolemia medications, and estrogen replacement therapy. It seems apparent that smoking avoidance by AAT-deficient nonsmokers and cessation among AAT-deficient smokers would be highly clinically effective and cost effective as primary preventive strategies.

The results presented here are somewhat conservative in calculating benefits for several reasons. First, it is likely that a large percentage of adults would be able to receive AAT replacement therapy at an annual cost less than $30,000, even when the cost of treating side effects is included. Also, it is likely that improvements in therapeutic administration and reductions in the cost of producing Prolastin will occur as new technologies are developed. Additionally, COPD medical costs and patient survival after the onset of symptoms are likely to continue their upward trends; both of these developments would enhance the cost-effectiveness of therapy designed to reduce or delay the probability of COPD incidence. Because of lack of data, we did not include any medical costs for mild stages of COPD disease. However, based on our sensitivity analysis, this last bias is probably quite small: a 40 percent increase in annual medical costs (or in missed medical costs) would only improve the estimated therapy benefits by 2–4 percent (Table 3). Finally, because of data and methodology limitations, we have ignored important positive and negative quality-of-life considerations, including those associated with routine IV therapy and with the pain and suffering of avoided disease.

More precise cost-effectiveness values could be calculated if we knew whether AAT replacement therapy is effective in improving patient outcome and, second, whether it merely halts progression of disease or improves lung function. Most of the anatomical changes associated with AAT deficiency are considered irreversible. We also need to know whether the use of AAT replacement therapy before clinical disease onset will prevent the development of emphysema in later life. If therapy is found to be efficacious, the timing of therapy initiation in relation to the course of COPD disease will also be of crucial clinical and economic significance. These issues can only be resolved by randomized, controlled, prospective double-blind studies. Such studies should antedate any widescale use of therapy or broad-based screening programs.

Randomized controlled clinical trials of AAT therapy designed to detect minimally significant clinical differences might inadvertently allow for ineffective economic results. The "outcome efficacy" survival benefits would have to be substantial for the intervention to be readily cost-justified. Such benefits could be detected in a relatively small clinical trial. Similar economic considerations would also have implications for other therapies and biomedical research issues. ☐

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TABLE A.1—Proportion of Persons Alive at Beginning of Interval Who Died during Interval

Age (years)	PiZ Males Smoking	PiZ Males Nonsmoking	US Males	PiZ Females Smoking	PiZ Females Nonsmoking	US Females
25–29	0.2200	0.1800	0.0083	0.0100	0.0030	0.0030
30–34	0.1893	0.0514	0.0095	0.1111	0.0039	0.0039
35–39	0.2555	0.0300	0.0118	0.1562	0.0055	0.0055
40–44	0.3254	0.0856	0.0166	0.1641	0.0087	0.0087
45–49	0.3696	0.0684	0.0254	0.4639	0.0568	0.0142
50–54	0.5714	0.1143	0.0409	0.1960	0.2009	0.0229
55–59	0.5045	0.1892	0.0651	0.6625	0.3066	0.0355
60–64	0.3571	0.1236	0.0985	0.5386	0.3553	0.0547
65–69	0.5645	0.3190	0.1429	0.2217	0.2162	0.0806
70–74	0.9821	0.2678	0.2135	0.4916	0.3489	0.1229
75–79	1.0000	0.6489	0.3044	0.8203	0.5742	0.1875
80–84	1.0000	1.0000	0.4315	1.0000	0.8081	0.2993
85+	1.0000	1.0000	1.0000	1.0000	1.0000	1.0000

SOURCE: C. Larsson and National Center for Health Statistics (see references[3] and [8]).

APPENDIX

CLYS cost-effectiveness estimates have been developed for a number of medical technologies and thus provide a convenient and widely accepted comparative perspective on this intervention relative to other well-known medical interventions.[10-12] CLYS values provide estimates of the net cost of saving one additional year of life. CLYS is measured as:

$$\text{CYLS} = \Delta C/\Delta Y. \tag{1}$$

The net change in COPD medical cost, ΔC, includes the present value of lifetime AAT replacement therapy costs plus costs of treating any therapeutic side effects, while subtracting out the average savings in COPD treatment costs resulting from reduced probability of disease. The increase in life years resulting from intervention, ΔY, is calculated as the average increase in patient life expectancy, plus a fraction of the average reduction in patient COPD disability days.

Emphysema patients report more days of restricted activity and work loss than average, even after age- and sex-adjustment. National Center for Health Statistics data on the annual age- and sex-specific number of restricted activity days associated with emphysema are 66.8, 0.1, and 28.7 for men age 21–40, 41–60, and 61 or older, respectively, and are 143.8, 0.6, and 26.2 for women 21–40, 41–60, and 61 or older, respectively.[27] Quality-adjusted life year (QALY) measurements are often used to weigh episodes of illness (in days) with 1 representing perfect health and 0 representing death.[12] Various methods have been developed to calculate QALY weights. We arbitrarily chose to weigh restricted activity days at 75 percent of healthy days in our base-case analysis. As shown in Table 2, since the average number of emphysema disability days is small relative to the average reduction in life expectancy, any weight between 0 and 1 has little impact on cost-effectiveness estimates. It should be pointed out, however, that CLYS estimates that disregard change in disability days would implicitly assume that a day of illness is equivalent to a day of health.

Change in survival was calculated by adjusting Larsson's age-, sex-, and smoking-status-adjusted five-year PiZ hazard rates by the ratio of age- and sex-adjusted US average population hazard rates to Swedish non-PiZ subject hazard rates (see Table A.1). Column one of Table 1 has projected this survival data to the reductions in life expectancy associated with AAT deficiency for age-, sex-, and smoking-status-adjusted US cohorts.

CLYS Estimates

The CLYS estimation equations can be described iteratively, just as they are calculated in a computer spreadsheet. Abstracting from sex and smoking status variations, let λ_t be the discrete mortality hazard at patient age t for intervals of length τ (e.g., $\lambda_t = 0.22$ in the five-year interval for smoking males age 25-29; Table A.1). Then L_t, the discounted life expectancy at age t obeys;

$$L_t = D_{\tau/2}\lambda_t\,(\tau/2 - \delta) + D_\tau[(1 - \lambda_t)\,(\tau d + L_{t+1})], \tag{2}$$

where D_τ and $D_{\tau/2}$ are the time discount factors for time intervals of length τ and $\tau/2$, respectively; δ is the estimated cumulative time lost due to disability, discounted forward from the median time of COPD diagnosis to the midpoint of the τ-year interval when death from COPD occurs, and d is a survivor discount factor to adjust the τ years of survival forward to the end of the interval. Equation (2) contains discrete-time approximation errors, since we are assuming that the average death occurs at the midpoint of the time interval, when it is actually more likely to occur somewhat later in the interval. Also, as in the US Vital Statistics tables,[28] all individuals over age 85 are combined into final sex-specific age categories. These simplifications do not substantially affect the results.

Let λ^p_t be the τ-year hazard for a (sex-, smoking-adjusted) PiZ individual, λ^e_t be the τ-year hazard associated with a therapeutic efficacy of 100e percent, and λ^n_t be the τ-year hazard for a normal individual. Then we define "outcome efficacy" as;

$$\lambda^e_t = \lambda^p_t - (\lambda^p_t - \lambda^n_t)e. \tag{3}$$

One hundred percent efficacy would mean normal survival and 0 percent efficacy would mean the same mortality hazard as an untreated PiZ individual. Using equation (2), we define L^e_t to be the estimated life expectancy with therapy and L^p_t to be life expectancy without therapy. ΔL^e_t, the change in life expectancy at age t associated with a therapeutic "outcome efficacy" of e, would then obey;

$$\Delta L^e_t = L^e_t - L^p_t. \tag{4}$$

In the absence of therapy, excess medical costs due to PiZ-related COPD would obey:

$$M_t = D_{\tau/2}[(\lambda^p_t - \lambda^n_t)\,C_1] + D_\tau[(1 - \lambda^p_t)\,M_{t+1}], \tag{5}$$

where C_1 are the medical costs for treating fatal COPD discounted forward to time of death. The base-case analysis assumes that an average COPD fatality would experience a course of illness requiring annual medical treatments costing $14,647 per year during an eight-year period prior to death. This is a conservative estimate of medical costs, since costs are excluded for PiZ-related disease insufficiently severe to cause death, and for COPD prior to illness severity meeting the criteria in the Strauss *et al*, Bergner *et al* and Burrows *et al* studies.[8,9,17] As detailed in the text above, these downward biases on our medical cost estimates do not have a large effect on the CLYS calculations.

Under therapy, the medical and intervention costs due to PiZ-related COPD would obey;

$$M^e_t = D_{\tau/2}[(\lambda^e_t - \lambda^n_t)\,(C_1 + C_2/2)] + D_\tau[(1 - \lambda^e_t)\,(C_2 d + M^e_{t+1})], \tag{6}$$

where C_2 are the τ-year therapeutic intervention costs, including treatment of side effects or adverse reactions, discounted to the midpoint of the interval. In our base-case analysis, these are assumed to be $30,000 per year, or $146,733 per five-year interval using a 5 percent discount rate.

ΔM^e_t, the change in medical costs at age t associated with a therapeutic "outcome efficacy" of e, would obey;

$$\Delta M^e_t = M^e_t - M_t. \tag{7}$$

Thus, our measure of CLYS associated with a therapy of "outcome efficacy" e is $\Delta M^e_t/\Delta L^e_t$.

Clinical Trial Power Calculations

The two-tailed power estimate $1 - \beta$ for clinical trial sample sizes with a significance level of $\alpha = .05$ is calculated using the binomial probability formula;

$$\beta = \Phi\{[N^{1/2}(P_0 - P_1)]/[|P_0(1 - P_0) + P_1(1 - P_1)|^{1/2}] - 1.96\}, \tag{8}$$

where $\Phi\{\cdot\}$ is the cumulative normal distribution function, N is the sample size per treatment arm, P_0 is the (five-year) control group mortality rate, and P_1 is the treatment group mortality rate.[23]

Population Screening Cost Estimates

The estimated screening costs were calculated for the US population. It was assumed that, excluding the young and the elderly, there were about 100 million adults who might be considered for screening. Of these, roughly 30 percent would be current or recent smokers, and 10 percent would be emphysema patients. Twenty percent would be geneal relatives of emphysema patients. Screening costs, including follow-up testing of initial positives were estimated at $110 per screenee. It was assumed that the full battery of screening tests would be performed for all geneal relatives (assumed to average 3.3) of identified AAT-deficient adults. As described in the text, we assumed a false positive rate of 1 percent.

Regular Source of Ambulatory Care and Access to Health Services

Rodney A. Hayward, MD, Annette M. Bernard, MD, MSPH, Howard E. Freeman, PhD, and Christopher R. Corey, MA

ABSTRACT

Background: To examine why people lack a regular source of ambulatory care (RSAC) and explore whether this commonly used access measure accurately identifies population subgroups at risk for barriers to continuity care.

Methods: Using data from a 1986 national telephone survey, we performed a content analysis of subjects' verbatim reports as to why they lacked an RSAC (n = 5,748).

Results: The 16.4 percent of respondents who lacked an RSAC gave the following reasons: 1) financial problems, 8 percent; 2) local resource inaccessibility, 5 percent; 3) not wanting a regular source of ambulatory care, 61 percent; and 4) transitory loss of their regular source of ambulatory care, 18 percent. However, some sociodemographic subgroups reported substantially more problems with access barriers, and these disparities were often not detected by the global measure, RSAC. The poor were not more likely than the non-poor to lack an RSAC (odds ratio [OR] = 0.8; 95% confidence interval, [0.6, 1.1]), but were much more likely to lack an RSAC for financial reasons (OR = 5.2 [2.6, 10.6]). Similarly, rural respondents were not more likely than urban dwellers to lack an RSAC, but were more likely to lack an RSAC because of local resource inaccessibility (OR = 5.8 [2.8, 11.9]).

Conclusions: We conclude that the global measure, RSAC, is not an accurate indicator of whether population subgroups have access barriers to obtaining a source of continuity care. (*Am J Public Health* 1990; 81:434–438)

Introduction

Substantial effort and funds are devoted to monitoring access to health services in the United States.[1–10] Whether or not people have a regular source of ambulatory care has been a traditional measure in surveys on access and use of health services.[4–7,9,11–12] It has been recognized that access problems, such as financial barriers and provider unavailability, are not the only reasons for not having a regular source of ambulatory care.[13–14] Yet this indicator is often used to compare the access to continuity of care for different sociodemographic subgroups. It is sometimes assumed that the absence of a regular source of ambulatory care is primarily accounted for by access barriers. Whether or not this characteristic identifies groups that are at high risk of not receiving continuity of care because of access barriers has not been critically evaluated.

To evaluate the usefulness of this indicator as a measure of health services access, we analyzed data from the 1986 Robert Wood Johnson Foundation's access to health care survey. The analyses had three main objectives: 1) to identify the reasons Americans lack a regular source of ambulatory care; 2) to examine whether or not such reasons vary by demographic subgroup; and 3) to determine if the absence of a regular source of ambulatory care identifies sociodemographic subgroups at high risk for not obtaining continuity of care because of access barriers.

Methods

This 1986 national telephone survey used random-digit dialing and had a response rate of 76 percent. The Waksberg screening procedure was used to minimize the number of nonworking numbers dialed.[15] Persons with chronic or serious medical illnesses were oversampled to allow for adequate evaluation of access problems of persons with major health problems. The results presented are weighted to produce estimates for the US population. Detailed descriptions of the weighted-probability sampling technique have been published previously,[5,8,10,16] and the interview and schedule data are available from the Inter-University Consortium for Political and Social Research at The University of Michigan. Weighted and unweighted sample sizes by social characteristics are presented in the appendix.

As part of the survey, each interviewee was asked, "Is there one person or place in particular you usually go to when you are sick or want advice about your health?" Those who answered "no" to this question were asked: "Many people do not have one particular place to get medical care. What is the reason that you do not have a regular doctor or place to go?" Interviewers were instructed to record the responses to this item verbatim.

The survey on access to health care sampled 10,130 Americans, but verbatim responses were stored for a subset of the

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Address reprint requests to Rodney A. Hayward, MD, Division of General Medicine, Department of Internal Medicine, University of Michigan Medical Center, 1500 E. Medical Center Drive, 3116 Taubman Center, Ann Arbor, MI 48109-0276. Dr. Bernard is also with that Department; Dr. Freeman and Mr. Corey are with the Department of Sociology, University of California, Los Angeles. This paper, submitted to the Journal December 11, 1989, was revised and accepted for publication August 13, 1990.

TABLE 1—Proportion of Respondents Lacking a Regular Source of Ambulatory Care (RSAC)

	No RSAC	
	n	%
All Respondents	5,748	16.4%
Ages (in yrs.)		
1–12	821	6.4%
13–22	591	19.9%
23–44	1,985	23.7%
45–64	1,318	14.0%
≥65	1,010	9.2%
Health Status		
Excellent/Good	4,833	16.9%
Fair/Poor	887	12.7%
Sex		
Male	2,404	20.9%
Female	3,329	12.2%
Household Income		
Poor	442	18.4%
Near Poor	492	15.8%
Non-poor	3,977	16.0%
Insurance Status		
Insured	4,965	14.8%
Uninsured	450	31.2%
Race or Ethnicity		
Black/Hispanic	695	20.7%
White/Asian	4,809	15.5%
Place of Residence		
Urban	4,813	17.7%
Rural	795	15.3%
Education		
Non-High School Graduate	358	17.6%
High School Graduate	4,337	16.1%
College Graduate	552	18.6%

interviewees; this subset consists of consecutive cases obtained by random-digit dialing (n = 5,748), and there are no substantive differences between this subset and the overall sample with respect to key variables (regular source of ambulatory care, age, sex, geographic region, income, ethnicity, or insurance status).

To aid in developing a classification system for categorizing verbatim responses, we asked 20 health care providers to list reasons that people lack a regular source of ambulatory care. This generated a list of 82 specific reasons. This list was then simplified into a five-category classification structure: 1) financial problems; 2) local resource inaccessibility; 3) did not want a regular source of ambulatory care (divided into four subgroups, described below); 4) transitory loss of their regular source of ambulatory care; and 5) no response or response not classifiable. Pilot testing of a sample of actual responses revealed high reliability, and reviewers did not feel that further revisions of the classification structure were needed.

Financial problems included being unable or unwilling to pay the expected cost of medical care. Local resource inaccessibility included those who reported an inadequate supply of health care providers in their area or who reported logistical problems with getting care (i.e., problems with parking, transportation, or contacting their health care provider).

Two investigators (RAH and AMB) independently reviewed and categorized all verbatim responses. Inter-rater agreement was 96 percent for classification into the five main categories and 94 percent for classification into categories and subcategories. Upon joint review, most disagreements were easily resolved, and when reviewers continued to differ, the response was designated as non-classifiable.

Analyses were performed on five dependent variables: 1) the proportion who did not have a regular source of ambulatory care, and the proportion who lacked a regular source of ambulatory care because of 2) financial problems; 3) local resource inaccessibility; 4) not wanting one; and 5) transitory loss of a regular source of ambulatory care. Each variable was broken down according to eight respondent characteristics: age, sex, self-reported health status, household income, ethnicity, education, medical insurance status, and urban vs. rural residence. For respondents age 22 or younger, main wage earner's education was used.

We also evaluated associations between having a regular source of ambulatory care and receipt of cervical and breast cancer screenings;[10] seeking a physician for serious medical symptoms;[17] and needing medical, surgical, or supportive medical care and not being able to obtain it.[5,8] We measured receipt of care for serious symptoms using a method developed by Shapiro et al.[17–18] Each respondent was asked serially about one of five serious symptoms: 1) shortness of breath with light exercise or light work: 2) chest pain while exercising; 3) loss of consciousness; 4) abnormal bleeding; and 5) weight loss of 4.5 kb. or more not caused by dieting. Subjects who reported one or more serious symptoms during the past month were reported as having sought care if they had contacted a physician for any of the symptoms.

Results

Overall, 17.0 percent of respondents reported not having a regular source of ambulatory care. However, analysis of verbatim responses found that 0.6 percent of these respondents had misunderstood the question and actually had a regular source of ambulatory care. Most errors were due to subjects thinking they were being asked if they saw their doctor regularly (i.e., "I have a family doctor, but I don't go in that often."). Correction of these errors did not alter any of the associations between a regular source of ambulatory care and the respondent characteristics. This paper presents results using corrected responses.

After correction for the errant responses, we found that 16.4 percent of all respondents lacked a regular source of ambulatory care (Table 1). Bivariate comparisons showed that those aged 13 to 44 were more likely than other age groups to lack a regular source of ambulatory care. Also, not having a regular source of ambulatory care was more common among the uninsured, those in excellent or good health, males, Hispanics and Blacks.

Table 2 shows the reasons respondents lacked a regular source of ambulatory care, as a percentage of those lacking a regular source of ambulatory care and as a percentage of the total sample. Of the 841 respondents who reported not having a regular source of ambulatory care, 8 percent reported not having one because of

TABLE 2—Reasons for Lacking a Regular Source of Ambulatory Care (RSAC)

Reason for Not Having an RSAC	Percent of Those Who Lack an RSAC (n = 841)	Percent of All Respondents (n = 5,748)
Financial problems	8	1.4
Local resource inaccessibility	5	0.8
Did not want an RSAC	61	9.9
No health problems	41	6.7
Use alternative health care	5	0.8
Don't like to go to the doctor	5	0.8
Prefer multiple sources of care	7	1.1
No explanation	3	0.5
Transitory loss of RSAC	18	2.9
No response/unintelligible	8	16.4%
Lack an RSAC	100%	16.4%

financial barriers, and 5 percent because of local resource inaccessibility. In contrast, 61 percent reported not wanting a regular source of ambulatory care; most of these reported not having perceived a need for care. In addition, 18 percent reported a transitory loss of their regular source of ambulatory care, and 8 percent had no answer available. Overall, 2.2 percent of all respondents reported lacking a regular source of ambulatory care because of financial barriers or local resource inaccessibility (Table 2).

To evaluate the relative associations between the respondents' sociodemographic characteristics and the dependent variables (no regular source of ambulatory care and the four reasons people reported not having a regular source of ambulatory care), we computed logistic regression models for each dependent variable (Table 3). Respondents aged 13–44 were more likely than those in other age groups not to have a regular source of ambulatory care. This difference was distributed among the four reasons for lacking a regular source of ambulatory care (Table 4). Those in excellent or good health were more likely to lack a regular source of ambulatory care, chiefly because of not wanting one. Men were more likely than women to lack a regular source of ambulatory care due to a greater likelihood of having financial problems, not wanting a regular source of ambulatory care, and transitory loss of their regular source of ambulatory care (Table 3). The uninsured were more likely than the insured to lack a regular source of ambulatory care, due to a greater likelihood of having financial problems, not wanting a regular source of ambulatory care, and transitory loss of their regular source of ambulatory care.

The poor were not found to be more likely than the non-poor to lack a regular source of ambulatory care, but were more likely to lack a regular source of ambulatory care for financial reasons (Table 3). Similarly, level of education and place of residency were not associated with having a regular source of care, however, those with less education were more likely to report a financial barrier to obtaining a regular source of ambulatory care, and rural respondents were much more likely . . ." Level of education was even less likely to be associated with having a regular source of ambulatory care, but those with less education were more likely to report a financial barrier to obtaining a regular source of ambulatory care (Table 3). Yet rural respondents were much more likely to lack a regular source of ambulatory care because of local resource inaccessibility (Table 3). In all these instances, the global measure, regular source of ambulatory care, failed to identify those with greater risk for having access barriers to a regular source of ambulatory care.

However, lacking a regular source of ambulatory care was a risk factor for not receiving recommended medical care (Table 4). Those who did not have a regular source of ambulatory care were less likely to have received breast and cervical cancer screening and to have seen a physician when they had a serious medical symptom, and they were more likely to report being unable to obtain medical or surgical care they needed during the past year. They were also less satisfied with the medical care they had received.

Discussion

Whether or not people have a regular source of ambulatory care has been a commonly used process measure of access and is used as an indicator of whether the respondent has a source of continuity of care. Having a regular source of ambulatory care is felt to be particularly important for those with chronic conditions as well as for maternity and child care. Whether having a source of continuity of care improves efficiency or promotes better quality of care has not been well studied, and our results do not shed light on this issue. However, we did find that lacking a regular source of ambulatory care identifies a group that is less likely to receive recommended health services and that is less satisfied.

The presence or absence of a regular source of ambulatory care is also used to evaluate different population subgroups' access to obtaining a source of continuity of care. In this study, we found that the global measure, regular source of ambulatory care, could not be relied upon for this purpose. First, we found that most people who lack a regular source of ambulatory care do not report access barriers as the reason. Only 13 percent of those who lacked a regular source of ambulatory care reported problems with finances or availability of local health care providers, whereas 61 percent reported not wanting one. Although each study employed slightly different classification structures, our findings of why people lack a regular source of ambulatory care are consistent with previous reports.[13–14]

Second, in most instances, regular source of ambulatory care was a poor indicator of which groups were more likely to have access barriers to obtaining continuity of care. For instance, rural respondents were no more likely than urban respondents to lack a regular source of ambulatory care; however, a rural residence was the only major risk factor for having problems with local resource inaccessibility. Therefore, relying upon the global measure would be misleading in terms of the importance of access problems for rural residents in obtaining a source of continuity of care. Similarly, of the three variables most strongly associated with lacking a regular source of ambulatory care due to financial problems, two (household income and level of education) were not associated with the global measure, regular source of ambulatory care.

This study has several limitations. Telephone surveys bypass the 5 percent to 10 percent of the population who do not have telephones, resulting in underrepresentation of the poor, elderly, and uninsured,[19–21] groups that are particularly at risk for access problems. Still, supplemental face-to-face interviews con-

TABLE 3—Odds Ratios for Lacking a Regular Source of Ambulatory Care (RSAC) Controlling for Other Factors in Logistic Regression Models (n = 5,748)

	Odds Ratio (95% Confidence Intervals)				
	No RSAC	No RSAC because of financial problems	No RSAC because of local resource inaccessibility	Did not want an RSAC	Transitory loss of RSAC
Age (yr)					
1–12	0.2 (0.1, 0.3)	0.6 (0.2, 1.8)	0.1 (0.1, 0.3)	0.1 (0.1, 0.2)	0.3 (0.1, 0.7)
13–22	0.8 (0.6, 1.1)	0.6 (0.2, 106)	1.2 (0.4, 3.6)	0.6 (0.4, 1.0)	0.9 (0.5, 1.6)
23–44	1.0	1.0	1.0	1.0	1.0
45–64	0.5 (0.4, 0.6)	0.6 (0.3, 1.4)	0.8 (0.4, 1.9)	0.5 (0.4, 0.7)	0.5 (0.3, 0.8)
≥65	0.4 (0.3, 0.6)	0.4 (0.1, 1.0)	0.9 (0.3, 2.3)	0.4 (0.3, 0.6)	0.3 (0.1, 0.6)
Health Status					
Excellent/Good	1.4 (1.1, 1.9)	1.9 (0.8, 4.6)	0.5 (0.2, 1.2)	1.5 (1.3, 2.2)	1.2 (0.6, 2.2)
Fair/Poor	1.0	1.0	1.0	1.0	1.0
Sex					
Female	1.0	1.0	1.0	1.0	1.0
Male	1.7 (1.4, 2.0)	1.8 (1.0, 3.3)	1.0 (0.5, 2.0)	1.8 (1.5, 2.3)	1.4 (1.0, 2.0)
Household Income*					
Poor	0.8 (0.6, 1.1)	5.2 (2.6, 10.6)	0.9 (0.2, 3.1)	0.6 (0.4, 1.0)	1.5 (0.8, 2.9)
Near-Poor	0.9 (0.7, 1.2)	1.8 (0.7, 4.3)	0.8 (0.2, 2.7)	1.2 (0.8, 1.7)	1.0 (0.5, 2.0)
Non-poor	1.0	1.0	1.0	1.0	1.0
Insurance Status					
Insured	1.0	1.0	1.0	1.0	1.0
Uninsured	2.0 (1.5, 2.7)	3.0 (1.5, 6.0)	0.7 (0.2, 3.1)	1.8 (1.3, 2.6)	1.8 (1.1, 3.2)
Race or Ethnicity					
Black/Hispanic	0.9 (0.7, 1.2)	0.7 (0.3, 1.5)	0.9 (0.3, 3.1)	1.3 (1.0, 1.8)	0.6 (0.3, 1.2)
White/Asian	1.0	1.0	1.0	1.0	1.0
Place of Residence					
Urban	1.0	1.0	1.0	1.0	1.0
Rural	1.0 (0.8, 1.2)	0.7 (0.4, 1.4)	5.8 (2.8, 11.9)	0.7 (0.6, 0.9)	1.3 (0.9, 1.9)
Education					
Non-High School Graduate	0.9 (0.7, 1.1)	2.5 (1.3, 4.8)	0.4 (0.1, 1.2)	1.2 (0.9, 1.7)	0.8 (0.4, 1.5)
High School Graduate†	1.0	1.0	1.0	1.0	1.0
College Graduate	0.8 (0.6, 1.1)	0.5 (0.2, 0.9)	1.3 (0.6, 2.8)	1.2 (0.9, 1.5)	1.7 (1.1, 2.5)

*Poor is defined as household income <100 percent of the federal poverty level, near-poor is 100–150 percent of the poverty level, and non-poor is more than 150 percent of this level.

†Includes those with some college.

ducted in 1986 demonstrated, as others have found,[19–21] that inclusion of nontelephone households would affect percentage estimates only slightly. Nonetheless, small percentage differences in national surveys can represent millions of people. Furthermore, both survey techniques completely miss the homeless and those in institutions. In addition, the results are self-reported and are therefore dependent upon the reliability of respondents' recall. We are not able to determine whether population subgroups have different thresholds in reporting an access barrier. Also, 1.4 percent of respondents did not have a classifiable verbatim response. Although we found no pattern of one group being statistically more likely to have no available response, it may be that some of the weaker associations or trends might have been altered if responses had been complete. Finally, some of those who have a transitory loss of their regular source of ambulatory care or do not currently want one, may have unreported access barriers or may encounter access problems if and when they try to obtain one.

TABLE 4—Receipt of Recommended Medical Care and Patient Satisfaction among Those with and without a Regular Source of Ambulatory Care (RSAC)

	Have an RSAC %	Do not have an RSAC %	Differences (95% CI)
Pap smear within the past year*	56	46	10 (3, 17)
Breast examination by physician within the past year†	59	31	28 (18,37)
Mammogram within the past year**	22	8	14 (5, 23)
Doctor visit for serious symptom during the past month††	66	42	24 (10,27)
Needed medical or surgical care during past year, but was unable to get it	7	12	–5 (–7,–3)
Completely satisfied with last doctor visit	85	75	10 (5, 14)

*For women 20 years or older.

†For women 40 years or older.

**For women 50 years or older.

††Percentage of those respondents who had 1) shortness of breath with light exercise/work; 2) exertional chest pain; 3) loss of consciousness; 4) abnormal bleeding; or 5) unexplained weight loss during the past month, and who saw or spoke to their doctor about it.

We conclude that most people who do not have a regular source of ambulatory care report that they do not want one, and the global measure, regular source of ambulatory care, fails to identify hose sociodemographic subgroups who are at greater risk of having barriers to obtaining a source of continuity of care. Therefore, if we are truly interested in who is at risk for access barriers to obtaining continuity of care, the reason for lacking a regular source of ambulatory care must be evaluated. ☐

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APPENDIX—Weighted and Unweighted Sample Size for Those with and without a Regular Source of Ambulatory Care (RSAC)

	Weighted Sample Size*	Unweighted Sample Size
Total Sample	2,626	5,748
Has an RSAC	2,196	4,904
No RSAC	430	844
Financial problem	(36)	(65)
Resource inaccessibility	(21)	(45)
Don't want an RSAC	(262)	(507)
Transitory loss of RSAC	(76)	(159)
Unclassifiable	(35)	(68)
Age		
1–12 years	517	821
13–22 years	402	591
23–44 years	931	1,985
45–64 years	468	1,318
≥65	308	1,010
Health Status		
Excellent/Good	2,285	4,833
Fair/Poor	335	887
Sex		
Male	1,267	2,404
Female	1,359	3,329
Household Income		
Poor	250	442
Near Poor	235	492
Non-poor	1,825	3,977
Race or Ethnicity		
Black/Hispanic	435	695
White/Asian	2,156	4,809
Place of Residence		
Urban	1,813	3,840
Rural	795	1,890
Education		
Non-High School Graduate	358	945
High School Graduate	1,337	2,873
College Graduate	552	1,130

*Represents a conservative estimate of the effective sample size after accounting for design effects.

A Community-Based Study of the Use of Chiropractic Services

Paul G. Shekelle, MD, MPH, and Robert H. Brook, MD, ScD

ABSTRACT

Background: Little population-based information is available on the use of chiropractic services.

Methods: We analyzed data from the RAND Health Insurance Experiment (HIE), a community-based study of the use of health services. Insurance claim forms for all fee-for-service patients who completed the study were examined for all visits coded as being seen by a chiropractor. Services provided and patient-specified symptoms were taken from these forms. Population-based use rates were calculated for each HIE site. Use rates and services were calculated separately for first visits and repeat visits.

Results: There were 5,279 persons who contributed 19,021 person-years of exposure during the study; 395 different persons used 7,873 chiropractic services for a visit rate of 41 per 100 person-years and rate of use of 7.5 percent. Forty-two percent of all visits were for pain in the back. Spinal manipulation accounted for 61 percent of all services provided. Compared to non-users, users tended to be White, middle-aged, married, and high school educated. Seven-fold geographic variations in the use of chiropractic services were seen.

Conclusions: Chiropractors deliver a substantial amount of health care to the US population, and there are significant geographic variations in the rate and intensity of use of chiropractic services. (*Am J Public Health* 1991;81:439–442)

Introduction

Previous studies of the use of chiropractic services have used patient samples that were obtained from selected chiropractors' offices or clinics.[1–8] Such studies have shown that chiropractic users are more likely to be middle-aged, employed, and high school educated. No difference in gender has been seen. Most studies have shown use among races in proportion to their representation in the population, but one study reported significantly greater use by Whites.[2] Low back pain is the most common complaint of patients seeking chiropractic care (from 32 percent to 45 percent, depending on the study), and spinal manipulation is the most frequently provided service (from 76 percent to 92 percent of all services). Population-based use rates cannot be estimated from these studies. Therefore, we analyzed data from a prospective, large-scale, community-based population to answer the following questions: How many people use chiropractic services? What are their demographic characteristics? For what symptoms do they seek care? What services are provided? Are there geographic variations in use?

Methods

The RAND Health Insurance Experiment was designed to assess how varying the patients' share of costs for health services affected their use of services, their satisfaction with health care, the quality of their care, and the state of their health. Eligibility, enrollment, data collection, and principal results have been reported elsewhere.[9]

In brief, the study was conducted between 1974 and 1982 at six sites: Seattle, Washington; Dayton, Ohio; Fitchburg, Massachusetts; Franklin County, Massachusetts; Charleston, South Carolina; and Georgetown County, South Carolina. These sites were chosen to represent the four major census regions, both rural and urban areas, and differing levels of demand for health services. The population enrolled in the experiment, which was entirely civilian and under age 65, has been shown to be representative (across a wide range of variables) of the communities from which it was drawn and of the US population aged younger than 65.[10] Each enrollee was randomized to one of 14 fee-for-service insurance plans. All plans covered an identical wide range of services, including chiropractic care. Families were followed for three to five years.

Insurance claim forms of all fee-for-service patients who completed the study were examined for patient visits that had been coded as visits to a chiropractor. The symptoms we report here were the symptoms reported by the patients on their insurance claim forms as reasons for the visits. We took the services provided to be those for which reimbursement was sought by the chiropractor: visits, manipulations, X rays, laboratory tests, etc. Some patients and providers may have recorded more symptoms than others for equivalent conditions. We adjusted for this possibility in the following way. If multiple symptoms were recorded for a

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Address reprint requests to Paul G. Shekelle, MD, MPH, Department of Medicine, Division of General Internal Medicine and Health Services Research, University of California Los Angeles, CA 90024. Dr. Shekelle is with the UCLA Department of Medicine and the RAND Corporation, and Dr. Brook is with the UCLA Center for the Health Sciences and the RAND Health Sciences Program. This paper, submitted to the Journal October 19, 1989, was revised and accepted for publication June 4, 1990.

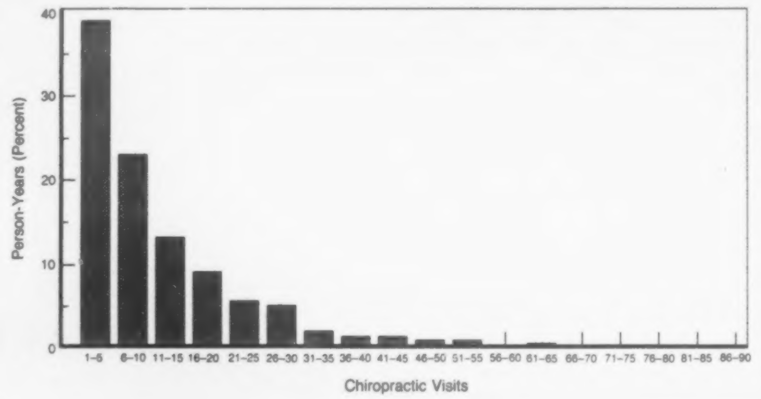

FIGURE 1—Number of Chiropractic Visits per Person-Year

TABLE 1—Comparison of Selected Demographic Characteristics of Chiropractic Users and Nonusers in the RAND Health Insurance Experiment

Characteristic	Chiropractic Users (%)	Nonusers (%)	Difference (95% Confidence Interval)
Male	49	48	1 (−4, 6)
White[a]	96	84	12 (9, 15)
Age			
<18	16	36	−20 (−24, −16)
18–50	70	50	20 (15, 25)
51+	14	13	1 (−3, 5)
Education[a]			
< High school	26	31	−5 (−10, 0)
High school graduate	47	36	11 (5, 17)
> High school	27	33	−6 (−11, −1)
Married[a]	74	45	29 (24, 34)
Working[a] (not homemaker)	67	64	3 (−2, 8)
Median family income[a] (1973 dollars)	$11,078	$11,192	114 (−666, 894)

[a] For adults age 18 or older.

single visit, each symptom was given a simple fractional weight so that for each visit the total symptom score summed to 1.0 (e.g., if two symptoms were recorded, each was given a weight of 0.5). Symptoms and services are presented here both as percentages and in terms of 100 person-years of exposure. A person-year was defined as a 12-month period beginning on date of enrollment (or anniversary date of enrollment) in the Health Insurance Experiment (i.e., not a calendar year). Population-based use rates were calculated for each Health Insurance Experiment site. Use rates and services provided were calculated separately for both the first patient visit and the repeat visits. First visits and repeat visits were defined by a specific code on the insurance claim forms to indicate whether the patient had seen this particular provider before. Because we could not calculate the number of visits on a per-episode basis, we calculated the number of chiropractic visits per person-year.

Results

The 5,279 people who completed the Health Insurance Experiment were enrolled for three or five years and represented a total of 19,021 person-years of exposure. During the study, 395 different persons (7.5 percent) made at least one visit to a chiropractor: 7,873 visits in all, or 41 chiropractic visits per 100 person-years. Repeat visits accounted for 82 percent of all visits. Only 1 percent of visits did not involve a face-to-face meeting between the patient and the chiropractor. Fewer than 1 percent of the visits were referred from another health care provider (either a physician or another chiropractor), seven percent of all visits were accident-related, and 1 percent were employment-related.

The 395 persons who saw chiropractors visited a chiropractor at least once during 683 person-years. In almost 40 percent of these person-years there were five or fewer chiropractic visits per year (Figure 1), and in 75 percent of them there were 15 or fewer visits per year. The median number of visits per year was seven and the mean was 11.5. In 13 person-years there were more than 50 visits per year, and these person-years contained 780 visits.

Table 1 compares the demographic characteristics of the chiropractic users with those of the nonusers. Chiropractic users were more likely to be White, aged 18–50, married, and high school graduates.

The most frequently given reason for a chiropractic visit (42 percent) was pain, swelling, or injury to the back region (Table 2). This symptom plus "back adjustment" accounted for 50 percent of visits overall, and this value varied from 41 percent to 69 percent, depending on the site.

Table 3 shows the services provided during the Health Insurance Experiment for the 395 chiropractic patients, by first visit and repeat visits. In the first chiropractic visit, manipulation (39 percent) accounted for the majority of services, but physical medicine visits (23 percent), office visits (19 percent), and X rays (17 percent) together accounted for almost 60 percent of the total services. Manipulation (66 percent) accounted for the majority of services provided for repeat visits, with physical medicine visits, office visits, and X rays accounting for 32 percent.

The frequency of chiropractic visits varied widely among the six Health Insurance Experiment sites (Table 4). The number of persons seeking care per 100 person-years ranged from 0.6 (Charleston) to 3.1 (Seattle). The number of services per 100 person-years on first visits and repeat visits varied from a low of 3.5 and 9.5, respectively, in Dayton and Charleston, to a high of 23.8 and 86.0, respectively, in Georgetown County, South Carolina.

Discussion

Our community-based study demonstrated a chiropractic visit rate of 41 per 100 person-years of exposure and that 7.5

TABLE 2—Patient-Specified Reasons for Chiropractic Visits

Reason for Visit[a]	Per 100 Person-years (N = 19,021 Person-years)	%
Pain, swelling, injury—back region	17.4	42.1
Pain, swelling, injury—face/neck	4.2	10.3
Headache	4.0	9.6
Back adjustment	3.4	8.2
Pain, swelling, injury—upper extremity	1.5	3.6
Pain, swelling, injury—lower extremity	1.2	2.8
Other symptoms referable to the digestive system	1.1	2.6
Progress visit, problem unspecified	1.0	2.3
Other examinations	1.0	2.2
General medical examination	0.8	1.8
Hip problems	0.6	1.4
Orthopedic foot problems	0.5	1.1
All others (75 different symptoms)	4.9	12.0
	41.6	100.0

[a]Includes all symptoms accounting for more than 1 percent of total visits.

TABLE 3—Services Provided by Chiropractors on First Visit and Repeat Visits per 100 Person-Years of Exposure

	First Visit		Repeat Visits	
Type of Service	Per 100 Person-years (N = 19,021 Person-years)	%	Per 100 Person-years (N = 19,021 Person-years)	%
Manipulation				
One area	4.1	38.9	25.8	66.1
Each additional area	<0.1	0.2	<0.1	0.2
Physical medicine visit	2.4	22.8	7.1	18.3
Various levels of office visit	2.1	19.3	4.2	10.7
Various X rays	1.8	16.8	1.2	3.2
Various laboratory tests	<0.1	1.0	0.2	0.5
All others	0.1	1.0	0.6	1.2
Total	4.2	100.0	45.5	100.0

percent of a community will use a chiropractor's service in a three- or five-year period. These numbers are somewhat lower than the 62 visits per 100 person-years and visit rate of 3.6 percent per year during the same time period that have been reported in the 1980 report to Congress on Manpower in the Chiropractic Profession.[4] Our study confirms previous observations that chiropractic users tend to be middle-aged and high school educated and that users do not differ from nonusers in their gender and income. We observed a greater use of chiropractic services by Whites, as also seen in one other survey.[2]

In previous studies, the average number of visits per user per year has varied between 5[6] and 18.[4] Our data, with a median of 7 and an average of 11.5, show that chiropractic use has a substantial tail to the right. In 13 person-years there were 50 or more visits per year. These 13 person-years (2 percent of the total person-years examined) contained 10 percent of the total number of visits. Results from the National Medical Ambulatory Care Survey between 1975 and 1980 showed that 7.7 visits per 100 person-years for back pain were made to physicians.[11] Our study showed that, during approximately the same time period, there were 17.4 visits per 100 person-years to chiropractors for the same complaint. Therefore chiropractors accounted for about twice as many visits for back pain as did physicians, as has been shown before.[12]

Symptoms of back pain accounted for 42 percent of patients visits. This finding is similar to 45 percent reported in Dade County, Florida;[5] 44 percent from Portland, Oregon[6]; and 32 percent from Tucson, Arizona.[8] This consistency from different geographic areas and time periods suggests that the symptoms for which patients seek chiropractic care vary little across the country.

Spinal manipulation, although still accounting for the majority of services provided, accounted for somewhat less than the 76 percent to 92 percent previously reported. In addition, there were differences in the types of services provided on the first visit and repeat visits. The typical first visit to a chiropractor appeared to consist of an office visit and/or physical medicine visit; most patients received an X ray and then manipulation. Repeat visits consisted of manipulation, with or without an additional office visit, and no further X rays. Because most chiropractic treatments involve a series of manipulations, it is not surprising that the number of manipulations far exceeded the number of office visits. As in previous studies, X rays and laboratory tests accounted for a relative minority of the total services.

First-visit use and repeat-visit use showed sevenfold and ninefold geographic variations, respectively, in chiropractic use in the Health Insurance Experiment. A previous study[4] had noted variations in chiropractic use, with greater use in the west and north-central regions of the US than in the South or Northeast. Our data cannot confirm these findings. In addition, Seattle and Charleston (both urban areas) had fourfold differences in use; Franklin County and Georgetown County (both rural) showed threefold differences in use. There was also substantial variation within a geographic area: the two southern sites exhibited the widest variation in use of chiropractic services.

The major limitations to this study are that its data are a decade old and that the elderly were excluded from the population. It has been estimated that the use of chiropractic services has been increasing during the past 20 years,[6] and the rate of use may be greater today than was seen in the RAND Health Insurance Experiment. A previous study of chiropractic use[2] reported that patients over age 65 accounted for 13 percent of the total number of chiropractic patients, and therefore we expect that excluding the elderly from our study will not qualitatively alter its conclusions. With these caveats, out study, because of its community-based design, provides an unbiased estimate of the use of chiropractic services, and documents that there are significant geographic vari-

TABLE 4—Use of Chiropractic Services for First Visit and Repeat Visits per 100 Person-Years, by Study Site

Site	Number Seeking Care per 100 Person-years	Services Used per 100 Person-years	
		First visit	Repeat visits
Dayton, OH	1.2	3.5	18.3
Seattle, WA	3.1	13.4	50.4
Fitchburg, MA	2.7	9.4	32.3
Franklin County, MA	1.9	7.1	30.5
Charleston, SC	0.6	5.7	9.5
Georgetown County, SC	2.8	23.8	86.0

ations in the rate and intensity of use of chiropractic services in the United States. ☐

Acknowledgments

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$7.5 Million Grant Program to Strengthen State Data Systems

The Robert Wood Johnson Foundation recently announced a $7.5 million grant program to be made available to state policy-makers to provide them with more accurate, up-to-date data about health care issues.

In the first phase of the program, up to 10 states will be awarded grants of up to $150,000 each to identify gaps in health policy and program information, and then to develop plans to fill them. Grant project directors will review information needs and set priorities, establish short- and long-term plans, and develop a specific proposal for enhancing existing data systems or for creating new ones to meet high priority information needs.

Each grant will be carried out by a lead government agency designated by the governor, and a working group composed of analysts from other state agencies, the legislature, state health program managers, health data agency representatives, private sector organizations, consumer groups, and others.

Based on the strength of the plans developed during the 18-month grant period, the RWJ Foundation will award as many as six four-year grants for amounts up to $1 million each; these grants will be used to develop and test the planned data system improvements, for data collection, analysis and dissemination.

The deadline for submitting proposals is August 1, 1991. Grant applicants must be state government agencies. An information conference for potential applicants is scheduled for May 1991.

Agencies wishing to apply for funds should contact: Ann P. Pomphrey, Program Assistant, Robert Wood Johnson Foundation, Route 1 and College Road East, P.O. Box 2316, Princeton, NJ 08543-2316. Technical assistance to the program will be provided by Ruth S. Hanft, PhD, research professor, Department of Health Services Management and Policy, George Washington University, Washington, DC.

Morbidity and Disability in Older Persons in the Years Prior to Death

37

Jack M. Guralnik, MD, PhD, Andrea Z. LaCroix, PhD, Laurence G. Branch, PhD, Stanislav V. Kasl, PhD, and Robert B. Wallace, MD

ABSTRACT

Background: A large proportion of the disease and disability which affects older persons occurs in the years just prior to death. Little prospective evidence is available which quantifies the burden of morbidity and disability during these years.

Methods: In three community-based cohorts of persons age 65 and older, chronic conditions and disability were evaluated for the three years prior to death in 531 persons who had three annual assessments and then died within one year of the third assessment. Number of chronic conditions, prevalence of disability in activities of daily living (ADLs), and prevalence of disability on a modified Rosow-Breslau scale were determined for these decedents and compared to 8821 members of the cohorts known to have survived.

Results: Prevalence rates of disease and disability increased during the follow-up for both decedents and survivors, with decedents generally having higher rates than survivors. Disability rates prior to death, but not the number of diseases, increased with increasing age at death. The odds ratio for disability in ADLs at any of the three assessments for decedents versus survivors ranged from 3.0 to 4.2 in the three communities. In each community the odds ratio for ADL disability was higher in women decedents versus survivors than in men decedents versus survivors.

Conclusions: These results have important implications for disability levels in future older populations in which death is projected to occur at increasingly higher ages. (*Am J Public Health* 1991;81:443–447)

Introduction

Prior to death, there is a decline in health status which may be quite variable. For the individual with a severe stroke who remains in a nursing home for several years, the state of poor health prior to death may be prolonged; for the person with metastatic cancer dying six months after diagnosis, duration of disability is moderate; and for the previously healthy individual who suffers sudden death, the state of poor health is virtually nonexistent. Just as the duration of poor health prior to death may be quite variable, its severity may also vary considerably. Increases in morbidity and disability prior to death have tremendous impact on individuals, their families, and the health care delivery system. With 71 percent of all deaths in the United States occurring among people age 65 and older,[1] it is of value to describe the systematic decreases in health status prior to death in this population.

The rising costs of caring for the older population have lead to a great deal of interest in the utilization and cost of health care in the years prior to death.[2] A study using Medicare data showed that 28 percent of all Medicare costs were incurred by the 6 percent of enrollees who would die within the next 12 months.[3] A study using data from the province of Manitoba was able to capture nearly all health care utilization in the years prior to death and found substantial increases in several measures of utilization with increasing proximity to death.[4]

Limitations are inherent in the use of health care utilization data to measure health status prior to death. Direct measures of disease and disability in older persons provide a more accurate description. There are a limited number of studies which examine disease or disability prior to death using either retrospective reporting of proxies[5,6] or prospective study designs.[7,8] In the present study, three representative populations of older persons were prospectively evaluated using self-report of common chronic diseases and two scales of physical functioning. The objectives of this analysis were to quantify the prevalence rates of disease and disability as death approached, compare these prevalence rates among decedents to rates in surviving members of the cohort, and examine how this process varies by gender and age at death. It was further hypothesized that disease and disability rates would increase to a greater extent in decedents than in survivors during the period of observation.

Methods

Study Populations

The data for these analyses are from the Established Populations for Epidemiologic Studies of the Elderly (EPESE),

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Address reprint requests to Jack M. Guralnik, MD, PhD, Epidemiology, Demography and Biometry Program, National Institute on Aging, 7550 Wisconsin Avenue, Federal Building, Rm 618, Bethesda, MD 20892. Dr. LaCroix is with the Group Health Cooperative of Puget Sound and the Department of Epidemiology at the University of Washington, Seattle; Dr. Branch is with Department of Socio-Medical Sciences and Community Medicine, Boston University School of Medicine; Dr. Kasl is with the Department of Epidemiology and Public Health, Yale University School of Medicine; Dr. Wallace is with the Department of Preventive Medicine and Environmental Health, University of Iowa. This paper, submitted to the Journal May 14, 1990, was revised and accepted for publication September 11, 1990.

TABLE 1—Number of EPESE Participants Surviving to Follow-up 3 and Dying between Follow-ups 2 and 3 by Site, Age Group, and Sex

Site and Age Group (years)	Survived to Follow-up 3		Died between Follow-ups 2 & 3	
	Men	Women	Men	Women
East Boston				
65–74	814	1333	46	37
75–84	314	583	30	47
85+	79	145	17	27
Total	1207	2061	93	111
Iowa				
65–74	746	1089	32	23
75–84	363	787	43	27
85+	80	181	12	21
Total	1189	2057	87	71
New Haven				
65–74	569	806	37	29
75–84	271	483	30	40
85+	49	129	16	17
Total	889	1418	83	86
Total All Sites	3285	5536	263	268

longitudinal studies of persons aged 65 and older. Participants included in these analyses were enumerated and interviewed in three communities in 1982 and early 1983. Details of study design and sampling strategies have been reported elsewhere.[9,10] Briefly, in East Boston, Massachusetts, and Iowa and Washington Counties, Iowa, full community surveys were conducted. A baseline response rate of 84 percent yielded 3,812 participants in East Boston, and in Iowa an 80 percent response rate yielded 3,673 participants. The New Haven target population was a stratified cluster sample of 3,420 eligible persons age 65 and older, drawn from three types of community dwellings: public elderly housing, where age and income are restricted; private elderly housing, where age is restricted; and private apartments and houses. There was oversampling of men and those living in elderly housing. A baseline response rate of 82 percent yielded 2,812 participants.

For the sake of clarity in understanding change in functional status and disease prior to death, these analyses were limited to a subset of participants for whom interview information was available for three consecutive years. The subjects of these analyses are those respondents from the three sites who participated in the baseline and first and second annual follow-up interviews and whose vital status was known for the next year of follow-up. Participants who died any time before the second annual follow-up were excluded from these analyses. Ascertainment of vital status at the end of the third follow-up year was nearly 100 percent at all sites.

Data Collection

The baseline interview was conducted in participants' homes and the first and second annual follow-up interviews were conducted by telephone or, if this was not possible, in person. A proxy interview was conducted for those unable to participate in the interview. Chronic conditions were assessed by asking respondents at baseline if they had ever been told by a doctor that they had a heart attack, stroke, cancer, diabetes, or hip fracture. At each follow-up interview, respondents were asked if they had been told by a doctor since the last interview that they had any of these conditions. Disability in activities of daily living (ADLs) was defined as self- or proxy-report of needing help from a person or special equipment in performing one or more of the following activities: walking across a small room, bathing, grooming, dressing, eating, transferring from bed to chair, or using the toilet.[11,12] Disability on a modified Rosow-Breslau scale of physical functioning was defined as the need for help in performing one or more of the following tasks: walking up and down stairs to the second floor, walking one-half mile, and doing heavy housework.[13] ADLs and Rosow-Breslau items were assessed at baseline and each annual follow-up.

Analytic Methods

Mean number of self-reported chronic conditions and prevalence of disability at baseline and at the first and second follow-ups were calculated separately for those dying in the one year period between follow-up 2 and follow-up 3 and for those surviving to follow-up three. The number of chronic conditions at each follow-up was calculated by determining the lifetime prevalence of each condition, taking into account baseline report as well as annual follow-up report of the occurrence of the condition in the previous year or years. Therefore, the number of chronic conditions for an individual could only stay the same or increase during the follow-up period. Individual reports of disability at each assessment were independent of other assessments.

Multiple logistic regression models were used to assess the odds of being disabled in any of the first three interviews for those dying between the second and third follow-ups compared to those surviving to follow-up 3.

Results

Table 1 shows the number of study participants who survived to follow-up 3 and who died between follow-ups 2 and 3 according to community, age group, and sex. There were 531 decedents and 8821 survivors in the communities by this definition. Those dying prior to follow-up 2 are not included in these numbers or in subsequent analyses.

In Figures 1, 2, and 3, the mean number of chronic conditions, the percent disabled in ADLs, and the percent disabled in Rosow-Breslau items and are shown for those dying between follow-ups 2 and 3 and those surviving to follow-up 3. Results are shown for baseline, follow-up 1 and follow-up 2 and are presented separately for each site, sex, and age group. For those dying between follow-up 2 and 3, the baseline data reflect morbidity and disability status from two to three years prior to death, the follow-up 1 data reflect status between one and two years prior to death, and the follow-up 2 data reflect status between zero and one year prior to death.

For nearly every site, sex, and age subgroup, the mean number of conditions and the proportion disabled was higher in those destined to die than in survivors. In most subgroups, these differences are evident even at baseline, which is two to three years prior to decedents' deaths. In general, mean number of conditions and disability prevalence rates increased as death approached. For those surviving to follow-up 3, there was also, for most subgroups, an increase in chronic conditions and disability from the baseline through the second follow-up interview. For some subgroups, the slope of the increase in de-

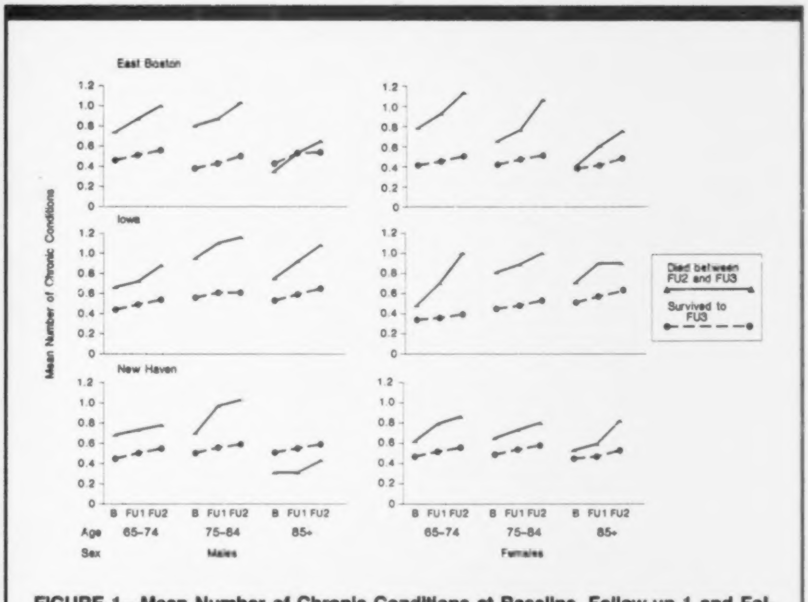

FIGURE 1—Mean Number of Chronic Conditions at Baseline, Follow-up 1 and Follow-up 2 According to Survival Status between Follow-up 2 and 3, Age Group and Sex; EPESE

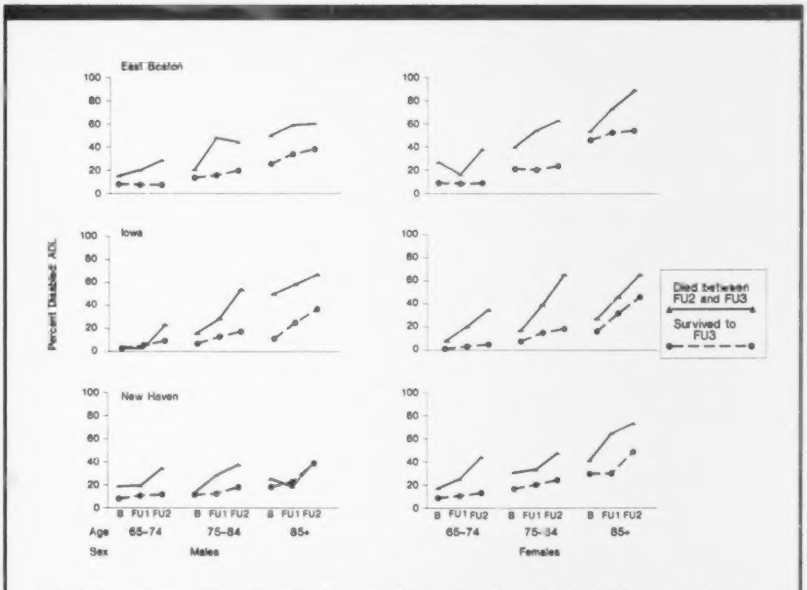

FIGURE 2—Percent Disabled in One or More Activities of Daily Living at Baseline, Follow-up 1 and Follow-up 2 According to Survival Status between Follow-up 2 and 3, Age Group and Sex; EPESE

cedents was clearly steeper than for the survivors, although this was not the case in all subgroups. In general, smaller differences were seen between decedents and survivors in the age group 85 years and older than for the two younger age groups.

Disability rates as measured by ADLs generally showed a substantial increase as death approached and were higher with increased age at death. In decedents aged 65–74 and 75-84 years at baseline, only about one-fourth or less report ADL disability at baseline, which was from two to three years prior to their deaths. At follow-up 2, within one year of death, the 65–74-year-old group had, in general, an ADL disability prevalence rate of 40 percent or less. In the 75–84-year-old group of decedents, up to two-thirds were disabled in ADLs in the year prior to death. Decedents aged 85 and older at baseline had the highest rates of ADL disability in the year prior to death (between 60 and 90 percent), with the exception of New Haven men.

Disability rates as assessed by the Rosow-Breslau items were higher in both decedents and survivors than corresponding rates of ADL disability. In the year prior to death, prevalence rates of disability, as measured by the Rosow-Breslau scale, were over 60 percent in nearly all subgroups, with those 85 and older having prevalence rates in excess of 80 percent. An increase in Rosow-Breslau disability rates with increased age at death was evident in most subgroups. Unlike disability measured by both instruments used here, mean number of chronic conditions prior to death did not show a consistent increase with age at death.

Table 2 demonstrates the magnitude of the differences in disability in ADLs in the baseline and first two follow-up interviews between those dying and those surviving. Odds ratios are presented for disability in any of these interviews for decedents compared to survivors. Overall, decedents were about three to four times as likely as survivors to report an ADL disability in either the baseline or first two follow-ups, after adjusting for age and sex. In men, decedents were 2.4 to 3.7 times as likely as survivors to report ADL disability. In women, the odds of disability in decedents was 3.7 to 5.2 times that of survivors. In each site, the odds ratio for disability for women who died versus survived was higher than the odds ratio for disability for men who died versus survived. The odds ratios for disability in decedents versus survivors by age group reveal no consistent trends with age across the three communities.

Discussion

Although the prevalence rates of disease and disability increase with age, there is no *a priori* reason to believe that within older populations those dying at a younger age should have less disability prior to death than those dying at older ages. These data reveal that those dying at the oldest ages generally have more disability, but not disease, prior to death than those dying at younger ages. As expected, the rates of disease and disability rise as death approaches, and those destined to die

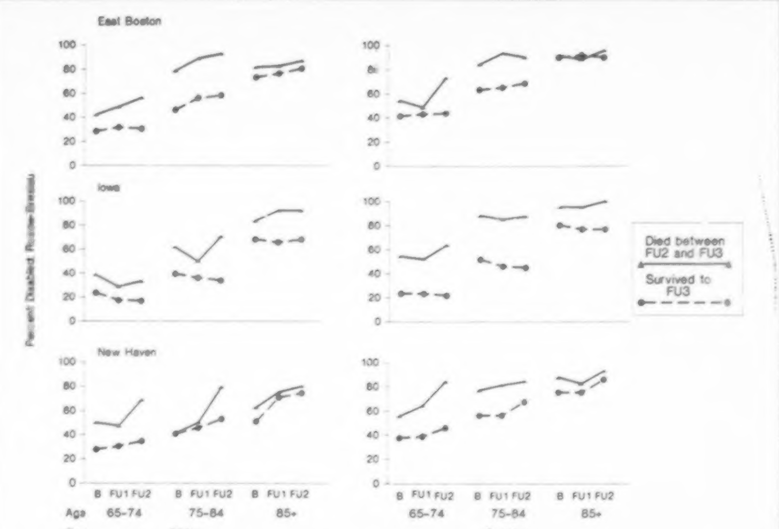

FIGURE 3—Percent Disabled in One or More Rosow-Breslau Items at Baseline, Follow-up 1 and Follow-up 2 According to Survival Status between Follow-up 2 and 3, Age Group and Sex; EPESE

have higher disease and disability rates than those of the same sex and age who survive.

In the conceptualization of these analyses, it was speculated that it might be possible to elucidate an even more refined relationship of morbidity and disability with death than we were able to demonstrate here. Consistent patterns were sought for differences in morbidity and disability during the three years of observation in decedents compared to survivors. Alternative patterns were postulated, including both groups starting out at the same disease or disability level and then diverging, the two groups starting out at different levels but remaining parallel, and the two groups starting out at different levels but still diverging during the years of observation. None of these patterns were observed in any consistent manner (Figures 1–3). This may be the result of methodologic shortcomings of this study or may reflect true variability in these relationships among different age groups, men and women, and the communities under study. A summary measure of ADL disability, that is, having any ADL disability in the three years prior to death, was employed to further investigate differences in sex and age relationships comparing decedents and survivors (Table 2). The odds of disability in decedents versus survivors showed no clear trend with age, but in each community the odds ratios for disability comparing decedents to survivors were higher in women than men.

A number of methodologic limitations must be considered in the interpretation of these results. The results in Figures 1–3 summarize percent disabled and mean number of conditions in a subgroup at a particular point in time, but changes for particular individuals as they approach death could differ considerably from these summary measures. It has been documented that change in functioning over time in individuals can be quite variable, with some reporting improvement, some reporting decline, and some reporting no change.[12,14] Research is now underway to analyze the multiple patterns of change which result from the repeated measures obtained in these longitudinal studies. These populations were all community-dwelling at baseline, although follow-up was completed for those entering an institution in subsequent years. Because rates of institutionalization increase dramatically with age, it is likely that the age gradient seen in disease and disability prior to death would be even steeper if all those in nursing homes three years prior to the deaths were included in the study. Although the cohorts under study are large, the number of deaths per year is not great and stratifying these deaths by age group and sex results in small sample sizes for certain subgroups. Larger numbers of deaths would have resulted in more accurate estimates of disease and disability prior to death and led to smoother curves for decedents in Figures 1, 2, and 3. The measures available to define morbidity and disability were fairly crude. The list of five chronic diseases was not comprehensive, being limited to those conditions queried at each interview. Disability status as estimated by both scales used here was dichotomized for ease of understanding and presentation. More refined scales, or perhaps the use of other measures which may be more sensitive to change over time, could lead to a finer understanding of the relationships investigated in this study. Finally, in comparing decedents to survivors there may be a "ceiling effect" which obscures the magnitude of differences in these two groups. This is seen, for instance, for the Rosow-Breslau scale for those aged 85 and older, in whom

TABLE 2—Odds Ratios (95% confidence intervals) for Reporting Disability in Activities of Daily Living at Any Interview (Baseline, Follow-up 1, or Follow-up 2) for Those Dying between Follow-up 2 and Follow-up 3 Compared to Those Surviving to Follow-up 3

	East Boston	Iowa	New Haven
Total Population*	4.0 (2.9, 5.6)	4.2 (3.0, 6.1)	3.0 (2.1, 4.2)
Men**	3.6 (2.2, 5.8)	3.7 (2.3, 6.0)	2.4 (1.5, 3.9)
Women**	4.4 (2.8, 6.9)	5.2 (3.0, 9.0)	3.7 (2.3, 6.1)
Age 65–74***	3.7 (2.3, 5.9)	3.3 (1.8, 6.1)	3.3 (1.9, 5.5)
75–84***	4.4 (2.5, 7.5)	5.5 (3.2, 9.5)	3.3 (2.4, 4.6)
85+***	4.4 (1.6, 11.8)	3.7 (1.5, 9.0)	1.9 (0.8, 4.2)

*Adjusted for sex and age.
**Adjusted for age.
***Adjusted for sex and age within 10 year age group.

the level of disability in survivors is so high that there may be little room for the decedents to appear more disabled (Figure 3).

There have been few prospective studies which have evaluated changes in disease and functional status prior to death. A number of studies have addressed terminal drop or terminal decline, but nearly all have focused on psychosocial or cognitive factors.[7] In a study which assessed declines in a number of domains, Palmore and Cleveland examined both terminal decline (defined as a steady, linear decrease prior to death) and terminal drop (defined as an accelerated decrease as death approaches) in 178 decedents in the Duke Longitudinal Study of Aging.[7] In a multivariate model, they found significant associations of both age and number of months prior to death with physical functioning. These results are in agreement with the findings presented here, as they can be interpreted to mean that, controlling for age at exam, there is decreasing physical functioning with increasing proximity to death, and, controlling for proximity to death, there is decreasing physical functioning with increasing age. They found no evidence of terminal drop.

Deeg and colleagues analyzed data for 88 residents of a home for the elderly in the Netherlands for whom four assessments were available in the year prior to death.[8] Declines were observed in physical, cognitive, and social functioning and morale as death was approached, and decedents had poorer scores in these domains than survivors. Model fitting revealed statistical evidence of an accelerated decline, or terminal drop, in physical functioning during this last year of life. In contrast to the present study, no age differences were found in physical functioning prior to death when comparing those younger than age 80 and those aged 80 and older. In another small study, there was no clear age gradient in the proportion of those judged clinically to be fully functional at 12 months, 6 months and 1 week before death.[15] These results suggest that there may be no age gradient in disability in the final year of life, but that this gradient may be present in the several years prior to death, as shown in this study. More research is necessary to answer this question definitively.

The current study has important implications for the older population now and in the future. Irrespective of the age at which an older individual dies, it is likely that a large proportion of the total years spent in the disabled state will be contributed by the years just prior to death. As the total population ages, preventive interventions, whether initiated at young or older ages, must have the ultimate goal of not only increasing longevity but of reducing disability. Any prevention efforts, at whatever age, to reduce total number of disabled person-years must address the issue of morbidity and disability prior to death, making this an important area for continued investigation. This study revealed that women decedents had proportionately higher rates of disability prior to death than men when compared to same-sex survivors, even after controlling for differences in age at death. This is important in an aging population in which, at each higher age, the female:male ratio increases and in which women have higher overall rates of disability than men.

Finally, the findings that the highest disability rates prior to death were found to be associated with death at the oldest ages can have grave implications for the total population burden of morbidity and disability in future years. If, as projected, life expectancy continues to increase,[16] the proportion of persons dying at the oldest ages will also increase. The total number of person-years of disability prior to death would thus increase, assuming age-specific rates of disability prior to death remain unchanged. Increased disability prior to death in future older populations, in which a large proportion of deaths will occur at the oldest ages, could make the quest for an overall compression of morbidity in the older population difficult, if not impossible. ☐

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The Yield of Cholesterol Screening in an Urban Black Community

Nancy Knowles Russell, PhD, Diane M. Becker, ScD, MPH, Cheryl P. Finney, MPH, and Hamilton Moses, MD

ABSTRACT

Background: While the distribution of cholesterol levels have been well studied in the general population, little is known about cholesterol and other cardiovascular disease risk factors in screenings held in an urban Black community. This study was designed to determine the yield of cholesterol screening in this community.

Methods: Screening took place in eight community sites. Serum total cholesterol was measured using a rapid capillary technique. Blood pressure was taken according to national guidelines and the average of two to three measurements were used. Standard interviews were used to determine the presence of cardiovascular risk factors by history.

Results: Of the 562 individuals screened, 44.9 percent had cholesterol levels requiring referral for follow-up care. Of those with total cholesterol ≥6.21 mmol/L, 66.4 percent were previously undetected and more than half also had blood pressure levels ≥140/90 mmHg on screening; 45 percent of all participants had blood pressure ≥140/90 on screenings. Of those with a history of elevated total cholesterol levels, none had levels below 5.17 mmol/L at the time of screening.

Conclusions: Multiple risk factors are highly prevalent in the urban Black community during cholesterol screening programs. Findings suggest the need for cholesterol programs incorporating blood pressure screening in the urban Black community. (*Am J Public Health* 1991; 81:448–451)

Introduction

Recommendations by the National Cholesterol Education Program (NCEP) and the availability of portable rapid cholesterol analyzers have facilitated the implementation of population-based mass cholesterol screening programs.[1–6] Recent studies indicate that Blacks are less likely than Whites to participate in voluntary mass screening programs, often held in integrated shopping malls or work sites.[7] Little is known about blood cholesterol patterns or the potential yield of cholesterol screening programs targeted specifically to an urban Black population in settings that minimize volunteer-bias. This is of particular importance to public health given that cardiovascular disease and stroke remain the leading causes of morbidity and mortality in the Black population.[8,9]

This study was designed to determine the yield of cholesterol screening and the prevalence of other coronary heart disease (CHD) risk factors in an urban Black population in East Baltimore.

Methods

Data were obtained from the first six months (June 1988 through November 1988) of the East Baltimore Community Heart Disease Prevention Project, a free collaborative cardiovascular risk screening and education program conducted by The Johns Hopkins Hospital, the churches in East Baltimore, and the community Mayor's Stations. The program is available to 71,292 residents in the 21 census tracts constituting East Baltimore.[10] Residing in these census tracts are 47,438 adults ≥21 years of age; 88 percent are Black; 53 percent are female; the median age is 20.5 years; 57 percent are high school graduates; 46 percent have income levels below the poverty limit; 30 percent of eligible adults are unemployed.[10]

Screening sites included five churches, one public housing project, and two major grocery markets in the catchment area. Screening was concentrated in churches, as 68 percent of Blacks in East Baltimore are reported to be active members of a church.[10] To minimize bias that may be associated with purely voluntary attendance at sites dedicated solely to screening, church screenings took place in concert with high attendance Sunday services, while screenings in grocery stores concentrated on Friday evenings and Saturday mornings.

Serum total cholesterol (TC) was measured using a rapid capillary (fingerstick) method and the Reflotron technology (Boeringer-Mannheim, Indianapolis). Participants were not fasting. Standard procedures for quality control recommended by the Laboratory Standardization Panel of the NCEP were followed at each screening site.[11] In a recent companion study of over 9,000 screening participants using the same cholesterol screening technology and personnel used in this study, this method was 95 percent sensitive for identifying TC levels ≥5.17

From the Johns Hopkins Medical Institutions, Baltimore, MD. Address reprint requests to Diane M. Becker, ScD, MPH, Johns Hopkins Community Heart Disease Prevention Programs, 1830 East Monument Street, 8037, Baltimore, MD 21205. Ms. Finney is also with the Heart Disease Prevention Program; Dr. Russell is with the School of Hygiene and Public Health; Dr. Moses is with the Office of the Vice President for Medical Affairs, Johns Hopkins Hospital. This paper, submitted to the Journal February 22, 1990, was revised and accepted for publication August 13, 1990.

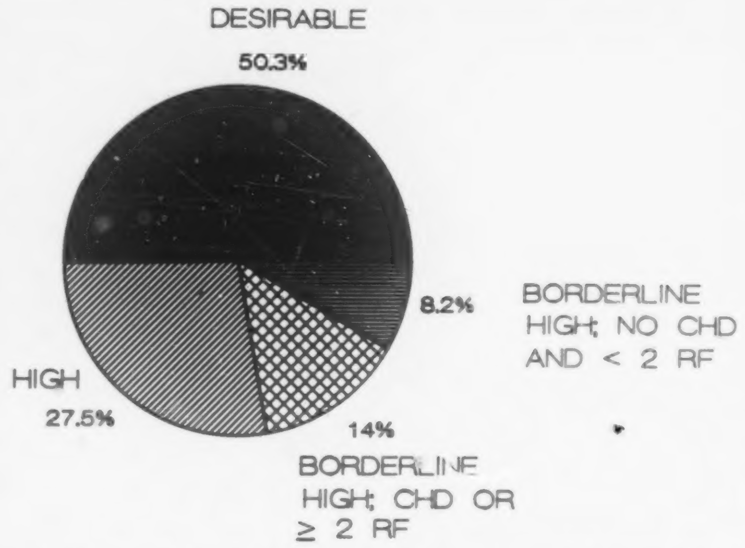

FIGURE 1—Overall Total Cholesterol Distribution according to National Cholesterol Education Program Adult Treatment Guidelines in Males (N = 186)

Borderline high = 5.17 to 6.18 mmol/L (200–239 mg/dl).
High = ≥6.21 mmol/L (240 mg/dl).
CHD = coronary heat disease.
RF = self-reported risk factors.

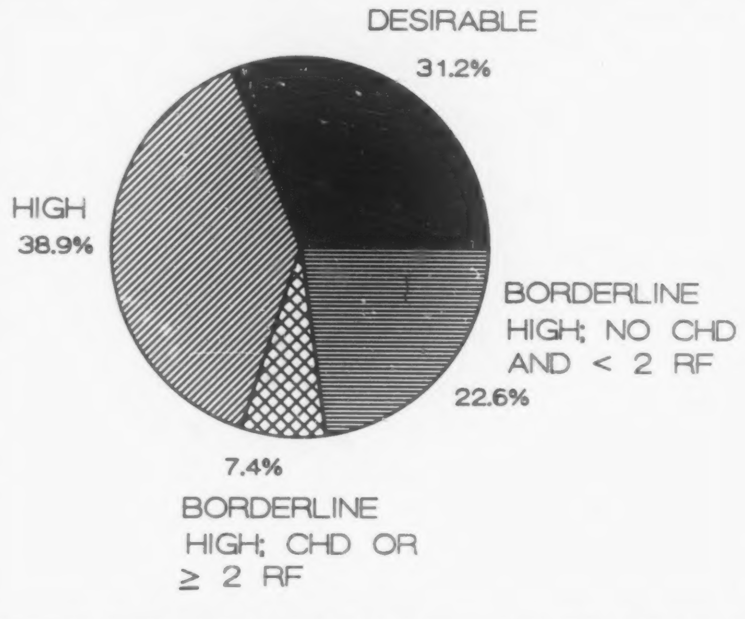

FIGURE 2—Overall Total Cholesterol Distribution according to National Cholesterol Education Program Adult Treatment Guidelines in Females (N = 366)

Borderline high = 5.17 to 6.18 mmol/L (200–239 mg/dl).
High = ≥6.21 mmol/L (240 mg/dl).
CHD = coronary heat disease.
RF = self-reported risk factors.

mmol/L (200 mg/dl), with the Johns Hopkins Lipid Research Clinics Laboratory method serving as the reference standard.[12]

Blood pressure measurements were taken with a mercury sphygmomanometer according to guidelines of the Joint Committee on High Blood Pressure IV (JNC IV).[13] The average of two measurements was used to characterize blood pressure during the screening for individuals with initial levels ≤120/80 mmHg; the average of three measurements was used for initial levels >120/80 mmHg.

Standardized personal interviews were conducted to determine the presence of CHD risk factors cited by the NCEP Adult Treatment Guidelines as factors that may lower the referral and treatment threshold for elevated cholesterol levels. Self-reported data were collected on age, height, body weight, smoking status, family history of premature CHD, medication use, and personal medical history of high cholesterol, hypertension, diabetes mellitus, CHD, and stroke. History of peripheral vascular disease proved difficult to assess by interview alone and was therefore not used. Accurate smoking history also proved difficult to obtain in the church setting, as many participating churches have explicit religious sanctions against smoking.

Results

The total number screened was 562; all were Black; the mean age was 50 ± 14.5 years for males and 51 ± 13.4 years for females; 66.7 percent were female; 11.8 percent had a personal history of CHD or stroke, and 10 percent had a history of diabetes. The mean TC was 5.35 ± 1.4[illegible] mmol/L for males and 5.86 ± 1.343 mmol/L for females. Among males, 41.5 percent had TC levels in the NCEP "high" and "borderline high" range with known CHD or two or more known risk factors (Figure 1). Among females, a total of 46.3 percent had TC levels classified as "high" and "borderline high" with existing CHD or two or more known risk factors (Figure 2). Thus, a total of 44.9 percent of all screening participants required referral to the medical care system for follow-up screening according to NCEP guidelines.

Data were also collected on prior awareness of cholesterol levels in 393 screening participants. The interview item on awareness was inadvertently omitted from printed instruments administered to the first 169 screenees, who did not differ

TABLE 1—Logistic Regression Analysis: Predictors of Total Cholesterol ≥6.21 mmol/L (N = 592)

Variable	Relative Odds	95% Confidence Interval:Odds
History of CHD or Stroke	1.17	1.01–1.35
Hypertension*	1.12	1.02–1.22
Female Gender	1.09	0.98–1.20
Age ≥60 years	1.05	0.38–1.15
History of Diabetes	1.03	0.90–1.18

*Hypertension = Blood pressure ≥140 mmHg or self-reported history of hypertension.

clinically or statistically in age, gender, TC levels, or co-morbidity from the next 393 participants. Of the 393 participants on whom the data were obtained, 56.3 percent reported never having had cholesterol levels checked before the screening. Of the 113 people identified as having TC levels ≥6.21 mmol/L on screening, 75 (66.4 percent) were not aware of high levels prior to the screening. More men (72 percent) than women (64 percent) were unaware of having elevated TC levels. Of those who reported having a known "high" TC level, 37 percent (23/61) had levels below 6.21 mmol/L at the time of this screening, but none were below 5.17 mmol/L; five people were taking a lipid-lowering medication.

More than half (54 percent) of those with TC levels ≥6.21 mmol/L also had hypertension, usually detected by screening rather than by history alone; 45 percent of screenees had blood pressure levels ≥140/90 mmHg on screening.

Logistic regression analysis was used to determine the relative importance of female gender, age ≥60 years, existing diabetes, a history of CHD or stroke, and the presence of hypertension in predicting "high" TC levels, ≥6.21 mmol/L (Table 1). Past history of CHD or stroke and hypertension, either by history or on screening, were the only principal independent predictors of having a "high" TC level. As body mass index (BMI) was calculated for only 126 participants, primarily because many individuals were reluctant to report weight or to be weighed in the mass screening setting, or did not know their height, this was not included in the overall regression analysis.

Discussion

Most prior screening studies conducted primarily in health care settings, work sites, and shopping centers in predominately White populations with a similar age distribution have yielded smaller percentages with NCEP "high" TC levels, generally in the 23 to 27 percent range[2,3,5,14] in contrast to the 35 percent found in the urban Black population in this study. However, in the second National Health and Nutrition Examination Survey (NHANES II) of a national probability sample of the civilian noninstitutionalized US population, a smaller percentage of Blacks (24 percent) were found to have NCEP "high" TC levels.[15] Furthermore, Whites were found to be more likely to have "high" or "borderline high" TC levels. This seeming discrepancy between the NHANES data and results of our study may be explained by the probable lower socioeconomic distribution of the community represented here.

Potential biases in the ascertainment of TC levels and subsequent NCEP classification may have resulted in underestimating the number of screening participants requiring referral to the medical care system. Specifically, we were unable to obtain accurate information for all individuals on smoking, obesity, and peripheral vascular disease in the screening settings. It is possible that larger numbers of individuals with "borderline high" TC levels had two or more NCEP-defined cardiovascular risk factors; this would place them at higher CHD risk and result in a medical care referral. Conversely, other biases in the sample may have resulted in overestimating the numbers requiring referral. Namely, women were differentially represented among participants. False positive misclassification bias may potentially have occurred in the studied population because women and Blacks are known to have higher levels of high density lipoprotein cholesterol, the lipoprotein subfraction that is potentially protective and that contributes to the total cholesterol level.[16] This potential source of bias may exist in this study as more women than men had "high" TC levels and slightly more women than men had levels that required referral to the medical care system according to NCEP Guidelines.

The association between hypertension and high cholesterol levels found in this study has also been reported recently in other population-based studies in Whites.[17] This link is of particular concern in the Black community given the known high prevalence of hypertension. Although the mechanism of this clustering of high cholesterol levels and high levels of blood pressure is not clear, replications of this finding suggest that the coexistence is not coincidental. The confluence of older age, diabetes, obesity, high sodium and fat diets, and sedentary life-style in urban Blacks may be important in the development of both hypertension and hypercholesterolemia.

This study, perhaps because of the nature of its venues, showed that women and older people were more likely to come to be screened. Cholesterol screening targeted specifically to Blacks in East Baltimore yielded a large number of people with previously undetected high cholesterol levels as well as large numbers of people with co-existing hypertension and hypercholesterolemia. These findings suggest the need for screening programs that address multiple rather than single risk factors in the Black community. Outreach strategies that extend beyond churches and food stores may be required to access men and younger people in urban Black community settings. In order to fulfill their ultimate mission to lower cardiovascular disease risk, cholesterol screening programs may need to focus strongly on the provision of health care follow-up, given the known dearth of chronic disease prevention services in inner-city Black communities.[18] Further, Blacks are known to utilize the health care system in unique ways that do not enable implementation of traditional health education or health promotion efforts.[19,20] Investigations of innovative programs that offer intensified on-site diet counseling and that provide at least first-stage verification of initial elevations of cholesterol may assist in diminishing referrals to the overburdened medical care systems in these communities.[19] ☐

Acknowledgments

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Call for Papers on Quality of Nursing Life: Taking Charge of Change

The Alberta Hospital Association, in cooperation with a number of nursing service, nursing education, professional, paraprofessional, employer, and labor organizations is sponsoring its second biannual forum in February 1992.

Quality Of Nursing Life: Taking Charge of Change is the theme of the forum to be held in Calgary, Alberta, Canada, February 5–7. The focus of the forum is to share innovative ideas for taking a leadership role in changes in health care.

Papers are being sought that present proven innovative ideas related to advancing positive change, including participation in decision making, dealing with the impact of economics, professional responsibility and ethics, promoting wellness for nurses, innovations in education, living with new information systems, promoting professional pride, or expanding roles of nurses.

A package providing details of abstract requirements is available from Education Services, Alberta Hospital Association, 10009-108 Street, Edmonton, Alberta, Canada T5J 3C5; (403) 498-8403, FAX (403) 498-8465. Deadline for receipt of abstracts is May 1.

Alcohol Drinking Patterns and Blood Pressure

Marcia Russell, PhD, M. Lynne Cooper, PhD, Michael R. Frone, PhD, and John W. Welte, PhD

ABSTRACT

Background: **Although blood pressure tends to increase with average alcohol consumption, little is known about the effects of drinking patterns on blood pressure. Therefore, the effects of average drinks per day and drinking pattern (defined as the independent and interactive effects of quantity and frequency) on blood pressure were compared.**

Methods: **Data were obtained from a random sample survey of 1,635 household residents in Erie County, New York. Alcohol-blood pressure relationships were examined using multiple regression analyses that controlled for the potentially confounding influence of 13 additional risk factors for elevated blood pressure.**

Results: **Consistent with prior research, a positive relationship was found between average drinks per day and diastolic and systolic blood pressure. Analyses examining the effect of drinking pattern indicated that drinking frequency had a positive effect on both diastolic and systolic blood pressure, whereas drinking quantity did not affect either. Furthermore, there was little evidence that the frequency-by-quantity interaction affected blood pressure.**

Conclusions: **Low average alcohol intake and low blood pressure were associated with infrequent drinking, rather than with frequent drinking of small amounts of alcohol. Results suggest that the standard practice of averaging alcohol consumption may obscure important effects of drinking frequency on health. (*Am J Public Health* 1991;81:452–457)**

Introduction

Despite the well documented relationship between average alcohol intake and elevated blood pressure,[1,2] we know relatively little about the effects of drinking patterns on blood pressure, where patterns are defined in terms of specific combinations of quantity drunk per drinking day and frequency of drinking days. Prior research has typically combined quantity and frequency measures of alcohol intake and expressed this information in terms of average consumption (average drinks per day, per week, or per month). However, such averaging may obscure potentially important differences in blood pressure associated with different drinking patterns. For example, although a person who drinks only two drinks every day of the week and one who has 14 drinks on Saturday night both average two drinks a day, blood pressures associated with these two patterns may not be the same.

To investigate the role of drinking pattern in the relationship of alcohol consumption and blood pressure, we examined the relationship between discrete combinations of quantity and frequency of drinking and average drinks per day; determined the effect of average drinks per day on blood pressure; determined the effect of drinking pattern on blood pressure by analyzing the independent and interactive effects of alcohol quantity and frequency; and contrasted the effects of drinking pattern and average drinks per day on blood pressure.

Methods

Sample

Data for the present analyses came from a representative adult household survey of stress, alcohol use, and hypertension conducted in Erie County, New York (N = 1,933). Respondents were identified using a stratified, three-stage probability sampling procedure designed to yield approximately equal numbers of Blacks and all others at three levels of education (less than high school, high school, at least some college). The first stage of sampling was the selection of US census blocks, stratified by race and education, and selected with probabilities proportional to size. In the second stage, housing units were selected with probabilities inversely proportional to the first stage selection probabilities. Finally, one adult (at least 19 years of age) was randomly selected from each household.

The overall sample completion rate was 78.3 percent, with the majority (84.5 percent) of noncompletions due to refusals. To provide population estimates in the present study, the sample was: poststratified by race, education, and sex (females were unintentionally oversampled) to match proportions reported in the 1980 census for Erie County; and weighted by the individual probability of selection.

Present analyses were conducted on a subset of 1,725 (weighted N = 1,630) respondents. Respondents over age 74 were eliminated (N = 75) because few drank alcohol, blood pressure measurements were missing for many, and doing so made our data more comparable to other studies of alcohol and hypertension. An additional 133 respondents who were

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From the Research Institute on Alcoholism, Buffalo, NY. Address reprint requests to Marcia Russell, PhD, Research Institute on Alcoholism, 1021 Main Street, Buffalo, NY 14203. This paper, submitted to the Journal July 25, 1989, was revised and accepted for publication September 27, 1990.

missing one or more of the study variables were deleted in order to ensure comparability across analyses. After exclusions, the weighted sample was 91 percent White, 54 percent female, and 63 percent married; 70 percent had completed at least a high school education. Mean age was 45 years (SD = 16.5). Forty-eight percent of the sample was working; of these, 51 percent held white collar jobs.

Procedures

Data were collected by a corps of 27 interviewers in the summer and fall of 1986. Interviewers received five days of intensive training on general and survey-specific interviewing techniques and three days of training on physical measurements, including blood pressure measurement. Interviews were conducted in respondents' homes using a highly structured interview schedule that included diet, smoking, and physical activity as well as medical history. The entire procedure took about 90 minutes; respondents were paid $25.

Measures

Blood Pressure: Standardized procedures using the American Heart Association protocol were used to measure systolic and diastolic blood pressures (SBP and DBP, respectively).[3] Blood pressure was assessed using mercury sphygmomanometers (Baumanometer, Model 300) calibrated for accuracy on a weekly basis. Respondents were asked not to drink alcohol during the interview. Properly fitted cuffs were placed on the right arm approximately 15 minutes before the end of the interview; respondents were asked to refrain from smoking, drinking coffee, or moving around the room until after their blood pressures were taken. Readings were made to the nearest 2 mmHg, and DBP was taken at the fifth Korotkoff phase (i.e., cessation of pulse sounds). Mean measures of SBP and DBP based on the last two of three readings were utilized in all analyses.

Alcohol Measures: Alcohol measures were adapted from the National Health and Leisure Time Survey.[4] Standard quantity and frequency questions (see Appendix) were asked on usual drinking during the past year. Average drinks per day was calculated by multiplying frequency of drinking by the number of drinks consumed on a typical day and dividing by 365. Quantity (number of drinks consumed on a typical drinking day—drinks per drinking day) was categorized as one or two drinks, three or four drinks, or five or more. Frequency (drinking days during the past year) was also categorized to yield groups of respondents who drank daily or almost daily, one to four times weekly, or less-than-weekly.

Nine drinking patterns were constructed by cross-tabulating the above quantity and frequency categories. They ranged from less-than-weekly drinkers who usually had one or two drinks per drinking day to those who drank daily or almost daily and usually had five or more drinks.

Covariates: The following potentially confounding factors were statistically controlled in all analyses: age (in years); race (1 = other than Black, 2 = Black); sex (1 = male, 2 = female); body mass index (weight in kg/height in meters, squared); marital status (0 = not married, 1 = married/living as married); education (in years); sodium intake (frequency of using salt at the table, in cooking, and eating salty snacks); calcium intake (milligrams per day based on the frequency of consuming calcium-rich foods, calcium supplements, and calcium-containing antacids); current cigarette use (number of cigarettes smoked per day); and two indices of frequency of physical activity.[5]

Also included as covariates were measures of family history of hypertension, self-reported hypertensive status, and whether or not antihypertensive medication was taken in the 24 hours prior to blood pressure measurement. Because a number of respondents were missing information on family history, two dummy variables were constructed, one to represent information concerning a positive family history of hypertension (family history: positive, coded as 0 = no, 1 = yes), and the other to represent missing information (family history: missing, coded as 0 = no, 1 = yes). Similarly, two dummy variables were used to represent information concerning respondents' self-report of their hypertensive status and current use of antihypertensive medication. The first (hypertension: no medication, coded as 0 = no, 1 = yes) represents individuals who reported that they had been told by a doctor or other health professional that they were hypertensive, but had not taken antihypertensive medication during the past 24 hours. The second dummy variable (hypertension: medication, coded as 0 = no, 1 = yes) represents individuals who reported that they were hypertensive and that they had taken antihypertensive medication during the past 24 hours.

Statistical Analysis

Three sets of statistical analyses were conducted. First, multiple regression was used to examine the relationship of average drinks per day to SBP and DBP, while controlling for covariates.[6]

Second, hierarchical multiple regression analysis was used to examine the independent and interactive effects of frequency and quantity on blood pressure.[6,7] More specifically, the hierarchical regression analysis proceeded in the following manner for both SBP and DBP: a block of covariates was first added to the regression equation in Step 1, followed by continuous measures of frequency and quantity in Step 2, which was followed by the interaction between frequency and quantity in Step 3. These hierarchical regression analyses did not contain abstainers because their inclusion would have led to an unbalanced design (i.e., a design with empty cells);[8] abstainers' scores on the quantity and frequency measures were both fixed at zero. Thus, respondents (i.e., abstainers) who had a zero on frequency could only have a zero on quantity, and vice versa. When a two-factor design is unbalanced, it is inappropriate to test for an interaction among its factors.[8] However, a two-way interaction can be tested if the design can be balanced by removing one or more levels of one or more factors.[8] In the present case, a balanced design was achieved by removing abstainers (i.e., the levels that represented zero frequency and zero quantity). Thus, we tested for a quantity-by-frequency interaction among drinkers only.

Third, to compare adjusted blood pressures in the nine drinking pattern groups with those of abstainers, we conducted two dummy variable regression analyses.[6] In each analysis, we examined the relationship of nine dummy variables to either SBP or DBP, while controlling for the covariates. Each dummy variable represented one of the nine drinking patterns with abstainers serving as the reference category. Therefore, in each regression equation, the unstandardized regression coefficient for a given dummy variable represented the mean difference in adjusted blood pressure for a specific drinking pattern compared to abstainers, and the constant represented the average adjusted blood pressure for abstainers.

As noted earlier, data were weighted to adjust means and regression coefficients for the complex survey design. However, standard errors have not been

TABLE 1—Average Drinks per Day (Mean ± Standard Deviation) by Drinking Patterns [weighted N = 1630*/100%]

	Frequency			
Quantity	Less than Weekly	Weekly	Daily	Total
Light (1–2 drinks)	.05 ± .05 [462/28%]	.48 ± .29 [237/15%]	1.37 ± .55 [58/4%]	.29 ± .43 [757/47%]
Moderate (3–4 drinks)	.18 ± .10 [163/10%]	1.12 ± .51 [212/13%]	2.91 ± .61 [39/2%]	.92 ± .89 [413/25%]
Heavy (5+ drinks)	.42 ± .20 [52/3%]	2.62 ± 1.59 [101/6%]	6.84 ± 1.46 [28/2%]	2.64 ± 2.43 [180/11%]
Total	.11 ± .13 [677/42%]	1.12 ± 1.09 [550/34%]	3.07 ± 2.29 [124/8%]	.79 ± 1.31 [1351/83%]

*Including abstainers: Weighted n = 279/17%.

TABLE 2—Regression Coefficients and Standard Errors of Multiple Regression of Systolic and Diastolic Blood Pressure on Average Drinks per Day and Covariates (N = 1630*)

	Systolic Blood Pressure		Diastolic Blood Pressure	
Variables	b	Standard Error	b	Standard Error
Age	.47	.03	.11	.02
Sex	−6.53	.78	−5.00	.51
Race	.06	1.29	1.19	.85
Body Mass	.34	.07	.26	.05
Marital Status	−2.12	.78	−.26	.51
Education	−.56	.15	−.09	.10
Sodium Intake	−.37	.20	−.38	.13
Calcium Intake	−.002	.001	−.003	.001
Family History: Positive	2.07	.87	1.14	.57
Family History: Missing	1.31	1.55	−.63	1.02
Hypertension: No Medication	8.78	1.02	5.73	.67
Hypertension: Medication	9.32	1.22	2.35	.80
Current Cigarette Use	.01	.03	−.005	.02
Activity	.46	.41	−.24	.27
Exercise	−.004	.02	−.01	.02
Average Drinks per Day	1.21	.32	.55	.21
Constant	108.13	4.03	71.59	2.64

*Abstainers included.

adjusted for design effects because the requisite statistical programs to do so were not available to us.

Results

Drinking Pattern

Mean values for average drinks per day associated with each pattern are presented in Table 1. As expected, marginal values for average drinks per day increased both as the frequency of drinking increased (r = .64) and as the usual quantity consumed on a typical drinking day increased (r = .56). Heavy drinking was more prevalent among daily drinkers than among those who drank less frequently. However, the magnitude of the relationship among drinkers is relatively modest (r = .23), which provides further justification for analyzing the independent and interactive effects of quantity and frequency on blood pressure.

Although quantity definitely influences average drinks per day, given the distribution of drinking in this household sample, frequency seems to play a more important role than quantity in determining which drinking patterns are associated with lower mean values of average drinks per day and which are associated with higher values. For example, mean averages of one drink or more per day were associated with moderate and heavy weekly drinking, and with daily drinking, irrespective of quantity; averages of less than one drink per day were associated with light, weekly drinking and less-than-weekly drinking, irrespective of quantity.

The most prevalent drinking patterns, accounting for 43 percent of the population, were characterized by light drinking less-than-weekly or weekly, both of which were associated with low average drinks per day. In contrast, only 8 percent of the population drank as often as daily or nearly every day.

Blood Pressure

Results of the multiple regression equations predicting SBP and DBP from average drinks per day are summarized in Table 2. Blood pressure had the expected relation to all covariates except race and sodium intake. Failure to observe higher average blood pressures among Blacks than Whites reflects their relatively small weighted numbers in these population-based analyses. Unweighted analyses are planned to investigate racial differences in the relation of blood pressure to drinking patterns. The negative relation of blood pressure to sodium intake reflects the lower discretionary salt intakes among individuals who had been told they were hypertensive, presumably in response to medical advice.

After covariates were taken into consideration, average drinks per day made a statistically significant contribution to the prediction of both SBP and DBP. The b values indicate that SBP was elevated an average of 1.21 mmHg, and DBP was elevated an average of 0.55 mmHg, for every increment of one drink in average drinks per day.

Drinking Pattern and Blood Pressure

Systolic Blood Pressure: The hierarchical regression analysis examining the independent and interactive effect of quantity and frequency on SBP revealed a significant positive relationship between frequency and SBP, but no significant effect for quantity and no significant quantity-by-frequency interaction (Table 3). The b value for frequency of drinking days during the past year indicated that SBP increased .021 mmHg with every additional day of drinking. Thus, SBP would be an average of 6.6 mmHg higher among those who drank daily than among those who drank only once a week (365 − 52 = 313 × .021 mmHg per drinking day per year = 6.6 mmHg).

TABLE 3—Regression Coefficients and Standard Errors of Hierarchical Multiple Regression of Systolic and Diastolic Blood Pressure on Drinking Frequency, Quantity, and Covariates among Drinkers (N = 1351*)

	Systolic Blood Pressure		Diastolic Blood Pressure	
Variables	b**	Standard Error	b**	Standard Error
Step 1				
Age	.47	.03	.13	.02
Sex	−6.88	.81	−5.80	.54
Race	.02	1.47	1.42	.99
Body Mass	.46	.08	.31	.05
Marital Status	−3.08	.83	−.93	.56
Education	−.48	.17	.09	.11
Sodium Intake	−.40	.21	−.33	.14
Calcium Intake	−.001	.001	−.002	.001
Family History: Positive	1.79	.91	1.29	.61
Family History: Missing	1.00	1.78	−1.20	1.19
Hypertension: No Medication	8.28	1.09	5.32	.73
Hypertension: Medication	11.19	1.35	2.23	.91
Current Cigarette Use	.02	.03	.01	.02
Activity	.88	.43	−.13	.29
Exercise	.02	.03	.01	.02
Step 2				
Quantity	.15	.19	−.011	.13
Frequency	.021	.004	.015	.003
Step 3				
Quantity-by-Frequency	−.001	.002	−.003	.001

*Abstainers omitted.

**Regression coefficients were taken from the step at which they were entered; therefore, each coefficient is controlled for all other variables entered at the same step and prior steps.

The dummy variable regression analysis comparing SBP among abstainers and drinkers found that compared to abstainers, mean SBP was significantly higher among light daily, moderate weekly, moderate daily, or heavy weekly drinkers (Table 4). These results paralleled the hierarchical regression results, indicating that relative to abstention, more frequent drinking was associated with increased SBP within each quantity category, whereas a tendency for SBP to increase with quantity was observed only among weekly drinkers.

Diastolic Blood Pressure: Overall, the main effects of the hierarchical regression analysis examining the independent and interactive effect of quantity and frequency on DBP were similar to those observed for SBP (Table 3). The b value for frequency indicated that DBP increased .015 mmHg with every additional day of drinking. Thus, DBP would be approximately 4.7 mmHg higher among daily drinkers compared to the individual who drank once a week.

In addition, there was a quantity-by-frequency interaction (Table 3). We explored the nature of the interaction by analyzing the effect of frequency among light, moderate, and heavy drinkers, and the effect of quantity among less-than-weekly, weekly, and daily drinkers. This set of follow-up analyses, however, failed to uncover a strong, readily interpretable, functional form for the quantity-by-frequency interaction. The fact that this interaction was weak and failed to replicate with systolic blood pressure indicates that it may be a chance finding.

The dummy variable regression analysis comparing DBP among abstainers and drinkers found that, compared to abstainers, mean DBP was significantly higher among light weekly, light daily, or moderate daily drinkers (Table 4). Again, these results supported the hierarchical regression results by indicating that relative to abstention, more frequent drinking increased DBP at every level of quantity, but that higher quantity did not increase DBP within the frequency categories.

Discussion

To summarize, our analyses revealed a significant increase in blood pressure, particularly SBP, with increasing average drinks per day. However, consideration of drinking patterns indicated that, among drinkers, this increase was influenced more by the frequency of drinking than by the quantity of alcohol consumed. Although a quantity-by-frequency interaction was observed for DBP, it was not observed for SBP, and its pattern had no ready explanation. Finally, relative to abstainers, respondents who drank less often than weekly tended to have lower blood pressures, and those who drank weekly or more often tended to have higher blood pressures.

As reviewed by Gleiberman and Harburg,[9] a number of epidemiological studies have reported that individuals having relatively low average alcohol intakes tend to have lower mean blood pressures than abstainers, exhibiting a U- or J-shaped alcohol-blood pressure relation. Although not statistically significant, this tendency was also observed in the present study, as indicated by the fact that blood pressure among abstainers tended to be higher than that among less-than-weekly drinkers, who had low average drinks per day. Studies are needed to determine whether very light or occasional alcohol use exerts a physiological protective effect or if it is associated with psychological or behavioral correlates related to lower blood pressure.[9]

Methods employed to measure and analyze alcohol-blood pressure relationships may influence its interpretation. Studies in alcoholic men[10] and pregnant women[11] found that self-reported frequency was more reliable than self-reported quantity. Thus, lower reliability may have contributed to our failure to observe a significant effect of quantity on blood pressure in the present study. Also, it is important to note that the prepotency of frequency in the relationship between alcohol use and blood pressure is obscured by employing measures of average alcohol intake. Analyses of drinking pattern, in which the independent and interactive effects of quantity and frequency are taken into consideration, clearly indicate that drinking patterns characterized by infrequent drinking are associated both with low blood pressures and relatively low average drinks per day. Conversely, daily drinking is associated both with high blood pressures and relatively high average drinks per day.

Data from the 1979 National Survey of US drinking practices indicate that light, daily drinking was reported by only 2 percent of the US population,[12] suggesting that relatively few people who report low average drinks per day are light, daily drinkers. Therefore, it may be misleading to attribute purported health benefits as-

TABLE 4—Regression Coefficients and Standard Errors of Multiple Regression of Systolic and Diastolic Blood Pressure on Drinking Patterns

	Systolic Blood Pressure		Diastolic Blood Pressure	
Variables	b	Standard Error	b	Standard Error
Age	.46	.03	.09	.02
Sex	−6.46	.79	−4.90	.52
Race	.12	1.29	1.23	.84
Body Mass	.36	.07	.29	.05
Marital Status	−2.14	.79	−.42	.51
Education	−.68	.15	−.19	.10
Sodium Intake	−.38	.20	−.39	.13
Calcium Intake	−.001	.001	−.002	.001
Family History: Positive	2.06	.87	1.03	.57
Family History: Missing	1.60	1.55	−.57	1.02
Hypertension: No Medication	8.52	1.02	5.63	.67
Hypertension: Medication	9.04	1.21	2.16	.79
Current Cigarette Use	.02	.03	−.002	.02
Activity	.45	.41	−.34	.27
Exercise	−.01	.02	−.01	.02
Light, Less-Than-Weekly	.87	1.13	−1.17	.74
Light, Weekly	2.10	1.36	2.24	.89
Light, Daily	8.55	2.17	6.87	1.42
Moderate, Less-Than-Weekly	−1.81	1.50	−.32	.98
Moderate, Weekly	3.46	1.43	.38	.93
Moderate, Daily	6.33	2.55	3.77	1.67
Heavy, Less-Than-Weekly	−3.50	2.35	−1.47	1.54
Heavy, Weekly	5.16	1.82	.81	1.19
Heavy, Daily	5.07	2.99	2.87	1.95
Constant (Abstainers)*	108.45	4.20	72.33	2.74

*Because abstainers served as the reference category, the constant represents the adjusted blood pressure for abstainers. The b weights for each drinking pattern (e.g., light, less-than-weekly) represent the magnitude of the difference in blood pressure for the drinking pattern compared to abstainers. Thus, the adjusted blood pressure for each drinking category is equal to the sum of the constant and its b weight.

sociated with low average drinks per day to light, daily drinking.

We believe that recognition of the importance of drinking frequency in determining the effects associated with average drinks per day is important for public health policy. The relatively low correlation between quantity and frequency among drinkers and the fact that few people in the US drink daily,[12] suggests that average drinks per day does not adequately represent actual drinking practices in the US. Accordingly, findings of health risks or benefits based on analyses employing average drinks per day cannot be translated directly into public health policy. The present data suggest that previously reported health benefits associated with low average drinks per day[2,9] are likely to be related to light, weekly drinking and infrequent drinking, rather than to light, daily drinking.

This is the first time that elevations in blood pressure have been attributed to drinking frequency, even in light drinkers, and it is important that these analyses be replicated. Assuming for the moment that they prove to be reliable, it is interesting to speculate about how frequent drinking might influence blood pressure. It has been hypothesized that individuals who recently drank alcohol may have elevated blood pressure because they are experiencing mild withdrawal symptoms or direct pressor effects of alcohol.[2] If this is so, it may be that individuals in our sample who reported drinking daily or nearly every day were more likely to be undergoing these effects. Criqui, *et al,*[13] have noted that repeated acute elevations in blood pressure related to frequent alcohol use are functionally equivalent to a chronic condition. Thus, elevations in blood pressure related to daily drinking could have significant implications for cardiovascular health and should not be dismissed as a short-term phenomenon.

Mean SBP was 6.6 mmHg higher and DBP was 4.7 mmHg higher among daily drinkers than among individuals who drank only once a week. Blood pressure elevations of this magnitude have both clinical and population significance. The association of DBP with stroke and with coronary heart disease (CHD) was recently investigated in nine major prospective observational studies, correcting for risk underestimates in previous reports related to "regression dilution" bias.[14] It was found that a long-term difference of 5–6 mmHg in usual DBP was associated with about 35–40 percent less stroke and 20–25 percent less CHD.[14] Data from the Multiple Risk Factor Intervention Trial[15] indicate that SBP is more strongly associated with CHD death than DBP, suggesting that the effect of frequent drinking on SBP may represent a greater CHD risk than is indicated by its effect on DBP.

The potential effect on health of relatively small reductions in mean blood pressure is illustrated by an overview of unconfounded randomized trials of antihypertensive drugs.[16] In 14 such trials, the mean treatment duration was five years, lowering mean DBP 5–6 mmHg; stroke was reduced by 46 percent, and CHD was reduced by 14 percent.

Within the DBP range covered by prospective observational studies (about 70–100 mmHg), there was no evidence of any "threshold" below which lower levels of DBP were not associated with lower risks of stroke and of CHD. Although increased risk may be small at lower DBP levels, the net effect is substantial when aggregated over the large numbers of people affected. Also, there are large numbers of individuals with moderately elevated blood pressures among whom relatively small increases might result in levels associated with diagnosed hypertension.

In sum, the effect of alcohol consumption on blood pressure has public health as well as clinical relevance. The present study illustrates how the standard practice of averaging alcohol consumption may obscure important effects of drinking pattern on health, and it is important that the study be replicated to determine the reliability of its findings. These results suggest that it may be useful to examine the independent and interactive effects of drinking quantity and frequency on other conditions thought to be influenced by alcohol intake, as well as in future epidemiologic studies of the alcohol-blood pressure relationship. ☐

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APPENDIX

Frequency: **Thinking back over the past 12 months, how often did you drink any kind of alcoholic beverage, that is, beer, wine, and liquor? Asked only of people who had reported drinking during the past year, respondents chose from eight categories: every day; almost every day; three or four times a week; two times a week; once a week; two or three times a month; once a month; or less than once a month.**

Quantity: **During the past 12 months, about how many drinks, on the average, would you have on a typical day when you drank? A drink is a 4 oz glass of wine, a 12 oz beer, or a drink containing 1 oz of liquor. Responses were open-ended.**

Caffeine Consumption During Pregnancy and Fetal Growth

Laura Fenster, PhD, Brenda Eskenazi, PhD, Gayle C. Windham, PhD, and Shanna H. Swan, PhD

ABSTRACT

Background: The purpose of this study was to examine the association between maternal caffeine consumption and low birthweight, intrauterine growth retardation, and prematurity, adjusting for multiple confounders.

Methods: Data obtained from birth certificates and interviews on 1,230 women with singleton live births were analyzed to evaluate the potential influence of caffeine consumption during the first trimester on fetal growth.

Results: The crude odds ratio for intrauterine growth retardation in infants of women reporting heavy caffeine consumption (>300 mg/day) was 3.86 (95% CI = 1.80, 8.40) which decreased to 2.90 (95% CI = 1.23, 6.87) after controlling for confounding factors. The adjusted odds ratio for low birthweight and heavy maternal caffeine consumption was also elevated (OR = 2.05; 95% CI = 0.86, 4.88). Women who reduced their caffeine intake from greater than 300 mg/day to less than that early in pregnancy had lower risks of delivering infants with either intrauterine growth retardation or low birthweight than women who continued to consume that amount. Preterm delivery appeared to be unrelated to caffeine consumption.

Conclusions: Taken together with studies reporting similar findings, these results suggest that heavy caffeine consumption increases the risk for fetal growth retardation. (*Am J Public Health* 1991;81:458–461)

Introduction

In 1980, the United States Food and Drug Administration warned pregnant women to restrict their intake of caffeine based on increased adverse reproductive outcomes in animals.[1] Since then, a number of epidemiologic studies have examined the association between caffeine intake and low birthweight (LBW) and a few have investigated its relationship to intrauterine growth retardation (IUGR). In a prospective study, Martin and Bracken found a more than two-fold risk of LBW in term deliveries associated with caffeine intake of 151–300 mg daily during the first trimester and a greater than four-fold risk associated with intake of >300 mg daily.[2] A case-control study which combined cases of LBW and IUGR found an increased risk associated with caffeine intake greater than 300 mg daily during the first trimester (OR = 2.94; 95% CI = 0.89, 9.65).[3] In addition, a retrospective study demonstrated a relationship between caffeine consumption of greater than 300 mg daily and reduced birthweight.[4] In contrast, a prospective[5] and a large retrospective[6] study found no relationship between caffeine or coffee consumption and reduced birthweight, respectively.

Methods

Participants and Questionnaires

Participants were women selected as controls in a case-control study of spontaneous abortion (SAB); a detailed description of subject selection and recruitment is presented elsewhere.[7] Birthweight was obtained from birth certificates. LBW was defined as birthweight less than 2,500 grams. Gestational age was calculated from the date of the last menstrual period (LMP) to the baby's birth date, as reported on interview. IUGR was defined as weight less than the tenth percentile of the mean weight at each gestational week using standards developed for California (all races and sexes combined).[8] Preterm delivery was defined as gestational age of less than 37 weeks.

The interview was conducted by telephone, on average, nine months after birth. For caffeinated beverage consumption questions (caffeinated coffee, tea and soft drinks), the participant was asked the amount consumed during the month before pregnancy, if it changed during pregnancy and, if so, when and to what amount. This information was used to calculate each participant's average consumption during the first trimester of pregnancy, based on her consumption each week. Caffeine consumption was calculated by summing the average amount of caffeinated coffee, tea, and soft drinks consumed, assuming a caffeine content of 107 mg/cup, 34 mg/cup, and 47 mg/can, respectively.[9]

The analyses were limited to singleton births (n = 1,252). Women who were missing information on birthweight (n = 1) or on caffeine consumption (n = 8) or whose calculated gestational age was 45 weeks or more (n = 13) were excluded from the analyses, leaving 1,230 singleton live births.

From the Reproductive Epidemiology Program, California Department of Health Services, Berkeley (Fenster, Windham, Swan); and the Departments of Maternal and Child Health and Epidemiology, University of California School of Public Health, Berkeley (Eskenazi). This paper, submitted to the Journal March 19, 1990, was revised and accepted for publication October 10, 1990.

Statistical Analyses

We examined maternal characteristics within four categories of caffeine consumption consistent with those from previous studies:[2–4] no caffeine; average caffeine use of 1–150 mg/day (light use); 151–300 mg/day (moderate use); and >300 per/day (heavy use). Analysis of variance was used to determine whether mean birthweights or mean gestational ages differed by category of caffeine use. Unadjusted odds ratios (OR), confidence intervals (CI) and chi-square trend tests were calculated for three levels of caffeine consumption with no consumption as the reference group.

Adjusted odds ratios were calculated using logistic regression with indicator variables for light, moderate, and heavy caffeine use. Caffeine was also modeled as a continuous variable. Since results were similar, categorical variables are presented for ease of comparison to other studies.

Confounding variables were included if they were strongly related to LBW or related to both LBW and caffeine use. The following maternal characteristics were entered into the LBW models: maternal age (≤ 34, >34); parity (0, ≥ 1); race (White, Hispanic, other); hypertension during pregnancy (yes, no); cigarettes smoked (0, 1–10 daily, ≥ 11 daily); and alcohol consumed (0, 1–3 drinks per week, ≥ 4 drinks per week). Separate models were run to test for an interaction between caffeine and tobacco use and between caffeine and alcohol consumption.

Reductions in caffeine consumption early in pregnancy were examined to determine whether these changes might affect fetal growth. Crude and adjusted odds ratios were calculated separately for women who did and did not reduce their intake from greater than an average of 300 mg/day to less than that amount within six weeks of their LMP. The reference group for these analyses was women with no caffeine consumption. Six weeks post-LMP was chosen as the time to evaluate the effect of reducing caffeine consumption because it approximates the earliest time that a woman realizes she is pregnant.

The strategy outlined above was followed for the univariate and multivariate analyses of caffeine consumption and IUGR or prematurity. The following covariates were entered in the models for IUGR: hypertension, cigarettes, and alcohol, as defined above. Variables included in the models for prematurity were: education (less than high school, high school graduate, some college, college graduate), race, hypertension, cigarettes and alcohol.

Results

Characteristics of Study Population

Table 1 presents population characteristics by four categories of caffeine consumption. The following variables were associated with increased caffeine consumption: multiparity, one or more therapeutic abortions, being employed, and use of cigarettes or alcohol. Ethnicity other than White or Hispanic was associated with decreased caffeine consumption, as was being more highly educated.

Association with LBW, IUGR

Overall, the risk factors for LBW were consistent with previous studies.[10] Nulliparous women over 34 years of age were at increased risk for delivering a LBW infant. Prior stillbirth, hypertension during the current pregnancy, smoking more than 10 cigarettes per day, or drinking four or more alcoholic drinks each week were associated with LBW. Race was also associated with LBW; Whites were least likely to deliver a LBW infant. Women who did not have insurance coverage for prenatal care (a surrogate measure for low socioeconomic status) were also at increased risk for LBW. IUGR was more common among women with hypertension during pregnancy, smoking more than 10 cigarettes per day, or drinking four or more alcoholic drinks per week.

Association of LBW and Caffeine

The mean birthweights for no, light, moderate, and heavy caffeine use were 3327 grams, 3311 grams, 3288 grams and 3170 grams respectively. Table 2 presents the unadjusted and adjusted odds ratios for LBW by caffeine consumption. A dose-response effect with increasing consumption of caffeine was observed (χ^2 trend = 4.03, p = 0.05). The crude risk for those with the heaviest level of caffeine consumption was twice that in the unexposed group. Adjusting for covariates resulted in a slightly lowered measure of association. Including interaction terms for cigarette smoking and caffeine consumption and for alcohol and caffeine consumption did not alter these estimates. Women who reduced their heavy caffeine consumption (within six weeks of their LMP) had a lower risk of delivering a LBW infant (adjusted OR = 0.65; 95% CI = 0.20, 2.11) than those women who continued to drink this amount (adjusted OR = 3.05; 95% CI = 1.09, 8.51).

Association of IUGR and Caffeine

The association of caffeine and IUGR increased with amount (χ^2 trend = 11.42, p = 0.0007), with a doubled risk for medium consumers and almost a four-fold risk for heavy consumers (Table 3). The adjusted OR for heavy caffeine consumption remained elevated, although decreased (OR = 2.90). Addition of interaction terms for cigarette smoking and caffeine consumption and for alcohol and caffeine consumption did not alter these estimates. Women who reduced their heavy caffeine consumption had a lower risk of IUGR (adjusted OR = 1.58; 95% CI = 0.56, 4.42) than women who continued this level of caffeine intake (adjusted OR = 3.74; 95% CI = 1.34, 10.41).

Association of Preterm Delivery and Caffeine

Mean gestational age varied little by caffeine use. Although the crude odds ratio for preterm delivery and heavy caffeine consumption was somewhat elevated (OR = 1.76), this estimate was decreased after multivariate adjustment (adjusted OR = 1.31; 95% CI = 0.63, 2.69). Women who continued their heavy caffeine consumption had a higher risk of preterm delivery (adjusted OR = 1.72; 95% CI = 0.70, 4.24) than women who reduced their intake (adjusted OR = 1.10; 95% CI = 0.50, 2.43).

Discussion

This study indicated elevated risks for IUGR and for LBW associated with heavy caffeine consumption which continued beyond the sixth week of pregnancy. Recall bias is an unlikely explanation of these findings since little relationship between preterm delivery and caffeine consumption was seen. Consistency of reporting of caffeinated beverage consumption was examined in a smaller case-control study of SAB,[11] in which caffeine consumption was asked as in this study. Substantial agreement for reports of caffeine consumption between interviews which were six months apart was found; approximately 77 percent of women's responses agreed to within one cup.[11] In addition, reporting of coffee consumption has recently been shown to have good reliability over a period of time which ranged from nine days to 6.5 years.[12]

TABLE 1—Distribution of Maternal Characteristics by Level of Caffeine Consumption during Pregnancy, Santa Clara County, California, 1986–87

		Daily Caffeine Consumption				
Variables*	Total	None	Light (1–150 mg)	Moderate (151–300 mg)	Heavy (>300 mg)	p-value
Subjects (N)	1230	378	609	165	78	
Subjects (%)	100	30.7	49.5	13.4	6.4	
Maternal Age (years)						
≤34	1072	30.6	50.5	13.1	5.8	p = 0.11
>34	158	31.7	43.0	15.2	10.1	
Parity						
0	507	31.2	53.3	12.2	3.3	p = 0.002
≥1	719	30.6	46.7	14.3	8.3	
Previous Spontaneous Abortions						
0	982	31.0	50.2	13.3	5.5	p = 0.14
≥1	244	30.3	46.3	13.9	9.4	
Previous Therapeutic Abortions						
0	969	32.9	48.8	12.7	5.6	p = 0.006
≥1	257	24.0	51.8	16.3	8.9	
Previous Stillbirths						
0	1197	30.9	49.5	13.4	6.2	p = 0.83
≥1	29	27.6	48.3	13.8	10.3	
Race						
Hispanic	320	28.1	50.3	15.0	6.6	p = 0.002
White	639	27.8	51.0	13.5	7.7	
Other	270	40.7	45.2	11.1	3.0	
Education						
<High School	202	27.7	49.0	16.8	6.5	p = 0.001
High School Graduate	357	23.3	51.5	15.7	9.5	
Some College	346	35.2	48.0	11.3	5.5	
College Graduate	325	36.0	49.2	11.1	3.7	
Marital Status						
Together	1123	31.0	49.8	13.2	6.0	p = 0.32
Apart	107	28.0	46.7	15.0	12.3	
Insurance Coverage						
Yes	1098	30.1	50.0	13.6	6.3	p = 0.60
No	132	35.6	45.5	12.1	6.8	
Employment						
Yes	870	27.9	51.4	14.0	6.7	p < 0.001
No	359	37.6	45.1	12.0	5.3	
Hypertension during Pregnancy						
Yes	45	33.3	37.8	20.0	8.9	p = 0.33
No	1183	30.6	50.0	13.2	6.2	
Cigarettes/day						
0	998	34.8	51.0	10.2	4.0	p < 0.001
1–10	146	18.5	42.5	26.7	12.3	
≥11	86	4.6	44.2	27.9	23.3	
Alcoholic Drinks/day						
0	566	39.2	45.9	20.8	4.2	p < 0.001
1–3	598	24.9	52.7	15.5	6.9	
≥4	62	9.7	53.2	17.7	19.4	

*Numbers for some variables do not total 1,230 because of missing data. Percentages for some variables do not total 100 because of rounding errors.

In these analyses, pre-pregnancy weight and pregnancy weight gain—maternal characteristics known to be associated with fetal growth—were not collected.[10] In Martin and Bracken's study, although caffeine users tended to gain less weight than nonusers, the caffeine risk estimates were only slightly reduced when weight gain was entered into the logistic models.[2] Birthweight and IUGR may also be influenced by exposures occurring during the second and third trimesters which were not ascertained. However, a case-control study found that average coffee and tea consumption in the second and third trimesters did not differ appreciably from that in the first trimester.[13]

There are two potential sources of exposure misclassification. We did not collect information on caffeine intake from caffeinated drug use, chocolate, or cocoa. This bias is unlikely to be substantial since for most pregnant women the largest source of caffeine intake is coffee.[4,14] Nevertheless, our measures of caffeine consumption were crude, given the substantial variation in caffeine content with method of beverage preparation and volume of serving.[9]

Our results suggest an effect of caffeine consumption of greater than 300 mg/daily on IUGR and LBW but not preterm delivery. These findings are consis-

TABLE 2—Odds Ratios for LBW and Caffeine Consumption, Santa Clara County, California, 1986–87

Caffeine Intake*	Case	Control	Crude OR	95% CI†	Adjusted OR††	95% CI†
0	27	351	1.00	—	1.00	—
1–150	34	575	0.77	(0.46, 1.29)	0.78	(0.45, 1.35)
151–300	14	151	1.21	(0.63, 2.38)	1.07	(0.51, 2.21)
>300	12	66	2.36	(1.17, 4.93)	2.05	(0.86, 4.88)
Total	87	1143				

*Average consumption during first trimester, including any changes in habit.

†Confidence Interval calculated by Cornfield's method (using Epistat).

††Adjusted by logistic regression for age, parity, race, hypertension during pregnancy and cigarette and alcohol consumption during pregnancy.

TABLE 3—Odds Ratios for IUGR and Caffeine Consumption, Santa Clara County, California, 1986–87

Caffeine Intake*	Case	Control	Crude OR	95% CI†	Adjusted OR††	95% CI†
0	17	361	1.00	—	1.00	—
1–150	41	568	1.53	(0.85, 2.68)	1.45	(0.80, 2.63)
151–300	14	151	1.97	(0.96, 4.08)	1.58	(0.73, 3.41)
>300	12	66	3.86	(1.80, 8.40)	2.90	(1.23, 6.87)
Total	84	1146				

*Average consumption during first trimester, including any changes in habit.

†Confidence Interval calculated by Cornfields Method (using Epistat).

††Adjusted by logistic regression for hypertension during pregnancy and cigarette and alcohol consumption.

tent with those of other studies[2–4] and are biologically plausible. Caffeine is readily absorbed from the digestive tract and freely crosses the placenta.[15,16] Caffeine increases intercellular cyclic 3′,5′-adenosine monophosphate (cyclic AMP), which may directly interfere with fetal development.[17] A dose of 200 mg of caffeine significantly decreases blood flow in placental villi,[18] probably through vasoconstriction.[19] Any reduction in uteroplacental circulation is strongly associated with decreased fetal growth.[10] IUGR[20] and LBW[10] are associated with increased risks of perinatal mortality and morbidity. Clinicians should counsel women early in pregnancy about reducing caffeine intake to below 300 mg daily. ☐

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Does Increased Detection Account for the Rising Incidence of Breast Cancer?

Jonathan M. Liff, PhD, John F. C. Sung, PhD, Wong-Ho Chow, PhD, Raymond S. Greenberg, MD, PhD, and W. Dana Flanders, MD, DSc

ABSTRACT

Background: The incidence of breast cancer has been increasing over time in the United States.

Methods: To determine the role of screening in this increase, trends in the incidence of *in situ* and invasive carcinoma of the breast were evaluated using records of the metropolitan Atlanta SEER program between 1979 and 1986. From a sample of records, evidence of symptoms and mammographic screening prior to diagnosis was recorded.

Results: The average annual age-adjusted incidence of invasive disease rose 29 percent among Whites and 41 percent among Blacks. Incidence increased in all age groups. A trend towards earlier detection of invasive disease was found. Asymptomatic tumors accounted for only 40 percent of the increased incidence among whites and 25 percent of the increased incidence among blacks, with mammography as the principal contributing procedure.

Conclusions: These data suggest that increased detection accounts for some but not all of the rising incidence of breast cancer in the United States. (*Am J Public Health* 1991;81:462–465)

Introduction

A rise in the incidence of breast cancer has been observed in the United States during the past decade.[1-5] In part, this phenomenon may be attributable to the increasing prevalence of suspected risk factors.[6-11] Alternatively, increased mammographic screening[12,13] could have resulted in more complete recognition of disease. This study was undertaken to determine whether the trends in breast cancer incidence could be explained by increased screening.

Methods

Data for this study were collected by a population-based cancer registry that has been described elsewhere.[14] Briefly, clinical and demographic data are collected on all new cases of cancer diagnosed among residents of five counties in metropolitan Atlanta. The majority of cases are identified from records at local hospitals, supplemented by case-finding from free-standing pathology laboratories, outpatient clinics, physicians' offices, and death certificates. The present study concerns all cases of *in situ* or invasive breast carcinoma first diagnosed among females between 1979 and 1986.

Race- and age-specific incidence rates of *in situ* and invasive breast cancer by stage were computed for each of four consecutive two-year time periods between 1979 and 1986, using population estimates derived by the US Bureau of the Census[15] and the National Cancer Institute. The incidence rates were age-adjusted to the 1970 US population 20 years of age and older, and approximate 95 percent confidence intervals (CIs) were calculated.[16] Age-adjusted incidence rate ratios with corresponding approximate 95 percent CIs[17] were calculated for each subsequent time period relative to 1979–80, and tests for trends over time were performed with Poisson regression.[18]

Probability samples of registry abstracts for patients with invasive disease (n = 100 Blacks, 100 Whites) in each of two time periods, 1979–80 and 1985–86, were also reviewed to characterize the process that led to cancer detection.

Results

The overall average annual age-adjusted incidence rates of both *in situ* and invasive breast cancer among Whites rose during this study (Figure 1). The climbing incidence of invasive malignancies, however, was confined to patients with localized disease at diagnosis. Among Blacks, no increase in the age-adjusted incidence rate of *in situ* breast carcinoma was observed until 1985–86 (Figure 2). The corresponding incidence of invasive breast cancer, in contrast, rose consistently over time. The increase in incidence occurred mainly in patients with localized disease, except for a climb in regional disease during the most recent time period.

Address reprint requests to Jonathan M. Liff, PhD, Associate Professor, Division of Epidemiology, Emory University School of Public Health, 1599 Clifton Road, NE, Atlanta, GA 30329. Dr. Flanders is Associate Professor, Dr. Chow is Assistant Professor, and Dr. Greenberg is Dean, all are with the Division of Epidemiology at Emory; Dr. Sung is Assistant Professor, Department of Community Health and Preventive Medicine, Morehouse School of Medicine. This paper, submitted to the Journal August 11, 1989, was revised and accepted for publication August 6, 1990.

For both races, incidence rates increased temporally in each of the age groups 20–49, 50–64, and 65 or over. Tables detailing the changes in rates over time will be provided upon request.

Age-adjusted incidence rates of *in situ* breast cancer more than doubled among Blacks from 1979–80 to 1985–86 (rate ratio [RR] = 2.3), and more than tripled among Whites (RR = 3.7). This increase was greatest for those aged 50 or older among Whites, and among those within the age group 50–64 among Blacks. For invasive breast malignancies, the temporal trends in incidence were more homogeneous across age groups. Overall, the rate ratios over time, with 1979–80 as the reference period, were greater for Blacks (45 percent increase) than Whites (26 percent increase).

Review of records for a sample of patients with invasive malignancies revealed that the proportion of patients who were asymptomatic or without a recorded symptom increased to 10 percent of both White and Black patients (Table 1). The majority of these women were diagnosed with a mammogram, either with or without an accompanying physical examination. The proportion of patients who were symptomatic but who were also diagnosed with a mammogram rose to 32 percent among Whites and 15 percent among Blacks (Table 1). Among Blacks, the increasing use of mammography was evident among patients with localized disease, but not among those with more extensive disease at diagnosis (Table 2). Among Whites, however, a substantial increase in the use of mammography was observed in both localized and nonlocalized disease categories.

Among Whites, about 2 percent of invasive cancers diagnosed in 1979–80 were asymptomatic (Table 1). Multiplication of this percentage by the age-adjusted invasive incidence rate for that time period (119.5 per 100,000) yielded an estimate of the age-adjusted rate of asymptomatic cancers (viz., 2.4 per 100,000). The corresponding estimate for Whites during 1985–86 was 15.1 per 100,000, or 40.2 percent of the total increase in breast cancer occurrence between these two time periods. The corresponding percentage for Black women was 25.4 percent. By similar reasoning, asymptomatic tumors reported as mammographically detected accounted for an estimated 20.1 percent of the increase among Whites and 12.8 percent among Blacks.

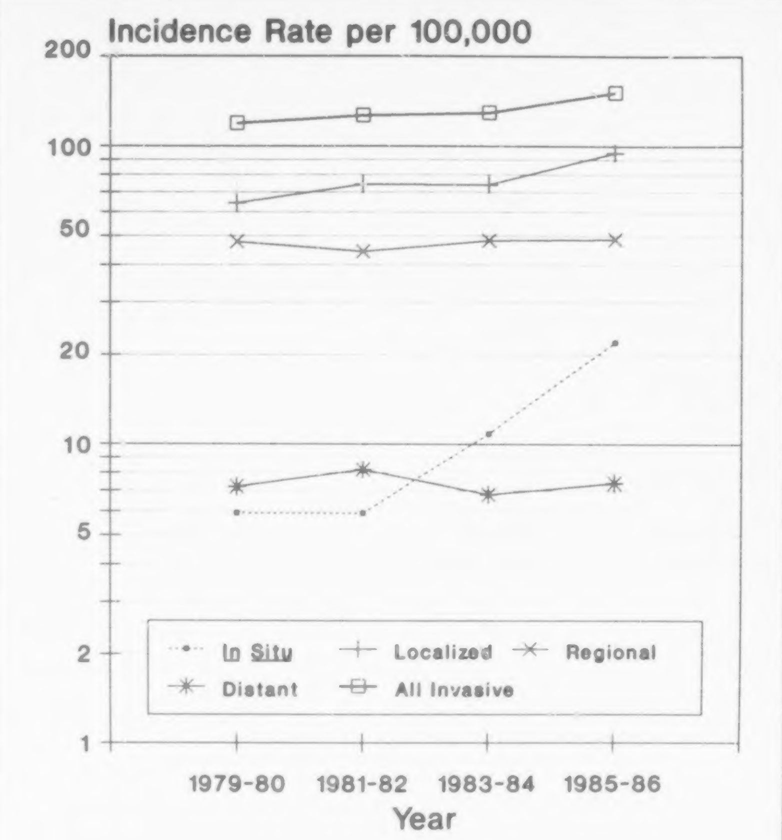

FIGURE 1—Trends over Time in Age-Adjusted Average Annual Incidence of Breast Carcinoma among White Women in Atlanta, by Stage.

TABLE 1—Percentage Distribution of Reported Detection Circumstances for Sampled Patients in Atlanta, Stratified by Race and Time Period of Diagnosis

	White		Black	
Detection circumstances	1979–80	1985–86	1979–80	1985–86
Symptoms Present				
Symptoms only	61	47	85	60
Physical examination	4	2	3	5
Mammography	16	29	4	15
Both examination and mammography	2	3	0	0
Other*	14	8	5	6
Symptoms Absent				
Physical examination	1	1	0	0
Mammography	1	5	0	4
Both examination and mammography	0	3	0	3
Other*	0	1	3	3
Symptom History Unknown	1	0	0	4
Total	100 (n = 100)	100 (n = 99†)	100 (n = 100)	100 (n = 99†)

*Detected during medical management of other conditions, such as back pain, gastroenteritis, diabetes mellitus, or benign breast disease.

†One medical record could not be located.

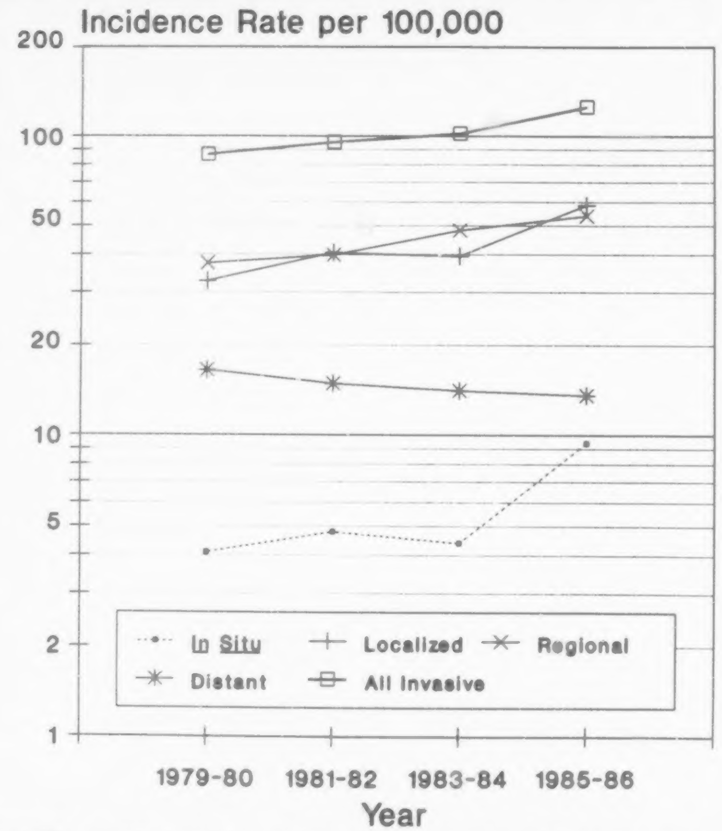

FIGURE 2—Trends over Time in Age-Adjusted Average Annual Incidence of Breast Carcinoma among Black Women in Atlanta, by Stage.

TABLE 2—Percentage Distribution of Reported Diagnostic Procedures and Stage for Sampled Patients in Atlanta, Stratified by Race and Time Period of Diagnosis

Diagnostic method	White 1979–80	White 1985–86*	Black 1979–80	Black 1985–86*
Localized Stage				
Physical examination	3	1	2	2
Mammography	13	25	1	15
Both examination and mammogram	2	2	0	3
Neither	35	32	32	28
Nonlocalized Stage				
Physical examination	2	[illegible]	1	3
Mammography	4	12	3	4
Both examination and mammogram	0	1	0	0
Neither	41	24	61	44
Total	100 (n = 100)	100 (n = 99*)	100 (n = 100)	100 (n = 99*)

*One medical record could not be located.

Discussion

The findings of increasing incidence rates of both *in situ* and invasive breast carcinoma in Atlanta are consistent with results from national surveillance data.[4] Also in accord with national trends,[4] among women with invasive malignancies, only those with localized disease demonstrated a substantial rise in incidence.

Although improvement in case ascertainment could produce an apparent rise in incidence, case ascertainment procedures did not change over time, and less than one percent of cases throughout this study were detected only from death certificates, suggesting virtually complete ascertainment at hospitals.

It has been shown that an increasing proportion of ductal carcinomas of the breast are detected by mammograms at the *in situ* stage.[13] In principle, mammography could lead to the detection of invasive cancers that would not be recognized otherwise. In the present study, however, mammographic detection of asymptomatic lesions accounted for only 20–40 percent of rising incidence among Whites and only 13–25 percent among Blacks.

The impact of mammography on earlier diagnosis of breast cancer was evident, however. The proportion of localized invasive carcinomas rose among Whites and Blacks, and the sampled abstracts revealed shifts toward mammographic diagnoses for localized disease among both races. These findings are in agreement with the results of previous studies, which have shown smaller tumor size, fewer lymph node metastases,[19] and earlier stage at diagnosis[20] among women with breast cancer detected by mammography screening.

Factors other than mammography could have contributed to the shift toward earlier stages of disease. In particular, heightened awareness of breast cancer[4,21] and increasing practice of breast self-examination[21,22] could result in earlier recognition of disease. Unfortunately, the medical records reviewed in this study did not include information on these health behaviors.

In summary, it does not appear that increased mammographic screening fully accounted for the rising incidence of breast cancer. Further work is needed to determine whether increasing prevalence of established risk factors or the emergence of new risk factors can be implicated. □

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NIMH Fifth Annual Research Conference on Mental Disorders

The National Institute of Mental Health is planning a research conference on the classification, recognition and treatment of mental disorders in general health care settings serving children, adolescents, and adult patients. The conference, planned for September 23–24 in the Washington, DC vicinity, will be the fifth annual conference on this topic.

Research presentations are encouraged for special sessions on: 1) diagnosis and care of mental disorders in *special populations* (rural, minority, child/adolescent) who are served by general medical caregivers; and 2) *alcohol and substance use disorders* in general medical settings. Other topics of interest include:

- the classification and accuracy of mental and substance use disorder diagnoses;
- the recognition and management of mental and emotional disorders found in the chronically medical ill;
- the severely mentally ill in general health settings;
- effectiveness of pharmacologic, psychosocial or consultation-liaison services;
- patterns and outcomes of referral of general medical patients to mental health specialists.

Research treatment settings may include ambulatory primary care settings, medical/surgical inpatient services, nursing homes, and emergency wards.

Abstracts, due May 30, will be reviewed by a scientific advisory committee and NIMH staff. Priority will be given to abstracts on new, completed research. Work in progress with an expected completion date before the conference, and original reviews of the research literature, may also be submitted.

For further information or to request an abstract form, contact: Junius J. Gonzales, MD, or Kathryn Magruder, MPH, PhD, NIMH, Fifth annual Primary Care Research Conference, Room 18C13, 5600 Fishers Lane, Rockville, MD 20857; (301) 443-1330/3364.

The Validity of Health Risk Appraisals for Coronary Heart Disease: Results from a Randomized Field Trial

Kevin W. Smith, MA, Sonja M. McKinlay, PhD, and John B. McKinlay, PhD

ABSTRACT

Background: **While health risk appraisals (HRAs) are becoming increasingly popular as tools for health assessment and health education, comparatively little is known about the accuracy of these risk estimates.**

Methods: **A field trial among 732 randomly selected adults ages 25 to 65 years was conducted to assess the validity of the risk scores produced by four widely used HRAs.**

Results: **Self-reported HRA risk scores for cigarette smoking and relative weight were generally accurate, but correlations between physiological measurements and scores for blood pressure, cholesterol, and physical activity were always lower than .51. Correlations between epidemiologic estimates of the probability of CHD death and HRA total risk scores ranged from .13 to .75; partial correlations adjusting for age, race, and gender ranged from .12 to .47.**

Conclusions: **The HRAs chosen for the field trial exhibited modest correlations with the CHD mortality risk predicted by the epidemiologic model. Mathematical errors made by respondents completing self-scored instruments and lack of knowledge regarding physiologic status decrease the accuracy of HRA risk estimates. (*Am J Public Health* 1991;81:466–470)**

Introduction

Health risk appraisals (HRAs) are designed to estimate a person's mortality or morbidity risk for various diseases based on such characteristics as medical history, blood pressure, and smoking habits. HRAs have become increasingly popular as health education tools. Recent estimates indicate that more than 200 organizations in the United States currently use HRAs to promote health awareness, and that between five and 15 million people have completed one of these instruments.[1] HRAs are also frequently included in work-site health promotion activities.[2] Despite the burgeoning popularity of HRAs, comparatively little is known about the validity of the risk estimates produced by them. Several recent reviews of HRA usage have stressed the need for further research on the accuracy of the predictions made by these instruments.[3-5]

In an earlier paper,[6] we examined the accuracy of the scoring systems employed by 41 different HRAs to estimate coronary heart disease (CHD) mortality risk. Correlations between HRA risk scores and epidemiologic estimates were found to be influenced by the sophistication of the HRA scoring system, the number of risk factors included in the instrument, the range of categories available for each risk factor, and the sensitivity of the scoring system to the effects of age and gender. Foxman and Edington[7] have also examined HRA accuracy with respect to mortality from all causes. These two studies based their assessments of HRA accuracy on physical examinations in the Framingham Heart Study and Tecumseh Community Health Study, respectively. Since the respondents who complete HRAs may or may not be aware of the appropriate values for risk factors, the use of actual physiological measurements in these studies is likely to have overstated the accuracy that would be obtained from self-reported risk scores.

This paper examines HRA validity based on self-reported data from a general population. Three questions are addressed: 1) How accurate are respondents' self-reports for HRA risk factors? 2) How accurate are HRA estimates of the risk of CHD mortality based on these self-reports? 3) To what extent is the validity of HRA risk scores affected by reporting errors and computational mistakes made by respondents?

Methods

A field trial of health risk appraisal instruments was conducted by the Cambridge Research Center of the American Institutes for Research in 1987. The trial was specifically designed to assess the reliability and validity of four representative HRAs in a general population. Respondents between the ages of 25 and 65 were identified through random digit dialing of 15 telephone exchanges in the Boston metropolitan area. Sampling and interviewing procedures for the field trial have been described elsewhere.[8]

During home visits, trained field technicians administered a standardized per-

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From the Cambridge Research Center, American Institutes for Research, Watertown, MA. Address reprint requests to Kevin W. Smith, Research Scientist, New England Research Institute, 9 Galen Street, Watertown, MA 02172. Dr. S. McKinlay is President and Dr. J. McKinlay is Vice President of NERI. This paper, submitted to the Journal March 3, 1989, was revised and accepted for publication August 13, 1990.

TABLE 1—Risk Score Ranges by HRA

	Arizona Heart Institute	CDC	Determine Your Medical Age	RISKO
Heart Attack Risk	0–50	2–21,969	—	0–27
Appraised Age	—	—	12.0–70.6	—
Smoking	0–10	.8–2.1	−6–24	−2–4
Relative Weight	0–3	.8–1.6	−6–24	−2–3
Blood Pressure	0–6	.4–1.1	0–12	−5–6
Cholesterol	—	.5–1.6	0–6	−3–3
Physical Activity	0–3	.6–2.5	−12–12	—

TABLE 2—Validation Measures for Risk Items and for Total Risk Scores

Self-Reported Risk Factors	Validation Measure
Blood Pressure	2nd Sphygmomanometer Reading (mmHg)
Total Serum Cholesterol	Laboratory Lipid Analysis (mg/dl)
Physical Activity	Harvard Alumni Activity Survey Scale (natural log of kcal expenditure for previous week)
Relative Weight	Quetelet Index (kg/m^2)
Smoking	Cigarettes per Day
Heart Attack Risk Appraised Age	NHANES I Epidemiologic Followup: CHD mortality risk estimate

sonal interview and one of the four HRAs to be evaluated by the project. Technicians were instructed not to help respondents complete HRA items or compute risk scores. The technician also measured blood pressure, height, and weight, and drew a venous blood sample for cholesterol analysis following the protocol established by the Centers for Disease Control (CDC). In-home physiological measurements have been shown to be as reliable and valid as those performed in a central location using standard equipment.[9]

Four widely used HRAs were chosen to represent the major types of computer-scored and self-scored instruments available in 1987: 1) Health Risk Appraisal (Centers for Disease Control), 2) The Heart Test (Arizona Heart Institute), 3) RISKO (American Heart Association), and 4) Determine Your Medical Age (Blue Cross/Blue Shield of New York). These were found to be among the most valid instruments within major HRA classifications in our previous work.[6] Since the field trial, the CDC instrument has been replaced by the Carter Center's "Healthier People" HRA.[10] The Carter Center's questionnaire and reporting format are similar to those of the earlier CDC version, but it uses an updated algorithm for assessing risk.

Each instrument requires a respondent to select categories that best describe personal habits, physiological status, and medical history. Point values or weights associated with these categories are then combined to produce an overall estimate of risk. In these four HRAs, total risk is summarized either by a heart-attack risk scale, with higher values representing greater risk, or by appraised age. Appraised age is the age at which the average mortality risk for all adults of the same gender and race is equal to the risk projected for a particular individual. If a person's appraised age is higher than his chronological age, that person has a greater than average chance of dying. The range of total risk scores and point values or weights for five major CHD risk items is shown in Table 1.

The CDC HRA, which requires computer processing, provides probability estimates for the 12 leading causes of death based on the Geller-Gesner method.[11] The Arizona Heart Institute's Heart Test and RISKO are self-scored instruments focusing on heart attack risk. In Determine Your Medical Age, appraised age is calculated by adjusting an individual's chronological age for medical history and lifestyle factors.

A standard psychometric definition for validity was adopted for this study; that is, that validity refers to the degree to which an instrument measures what it is intended to measure.[12,13] In order to assess the validity of a measure, some criterion or "gold standard" must be available for comparison. Table 2 shows the validation measures used in this study. Physiological values served as standards for three of the five CHD risk items. The standards for physical activity (kilocalorie expenditure for the previous week as calculated from the Harvard Alumni Activity Survey Scale)[14] and cigarette smoking were based on self-reports collected during personal interviews. HRA scores for heart attack risk or appraised age were compared to estimates from the NHANES I Epidemiologic Followup Study (NHEFS).[15] Separate NHEFS logistic models for men and women provide estimates of the probability of death due to coronary heart disease over a ten-year period based on a large (N = 4,013) representative sample of white adults.

Several statistical techniques can be employed to assess the extent of agreement or conformity between two measurements, including sensitivity, specificity, correlations, mean differences, and standard deviations for differences. Which technique is most appropriate remains a controversial issue.[16,17] However, the self-reported risk scores and the validation measures in this study cannot be directly compared because they were not measured on the same scale. Therefore, product-moment correlations between risk scores and validation measures were employed as the primary indicator of accuracy.[12]

As the criterion measure for CHD risk, the 10-year heart disease mortality probability for each respondent was calculated using the gender-specific NHEFS models. Physiological measures taken by field technicians, rather than self-reports, were used to make the logistic estimates as accurate as possible. According to the logistic models, a few respondents had mortality risks that were much higher than the risks for the great majority of cases. To prevent these outliers from distorting the correlation coefficients, the natural logarithm of the 10-year mortality risk per 100,000 persons was used in the analyses.

Three HRA total risk scores were computed for each respondent: 1) the score reported by the respondent, 2) the risk score corrected for mathematical errors, and 3) the score that would have been obtained on the basis of the blood pressure, cholesterol, height, and weight measurements taken by field technicians. The first of these scores does not apply to the CDC instrument, which is calculated by computer. Each of these three risk scores was then correlated with the log-

TABLE 3—Demographic Characteristics of Field Trial Participants and Nonparticipants*

	Participants		
Characteristics	Physiologic Status Measured at Baseline (N = 309)	Physiologic Status Measured at Follow-up (N = 423)	Nonparticipants (n = 544)
Age (years)	39.8 (10.7)	38.9 (10.6)	40.2 (11.1)
Percent female	59.9 (49.1)	56.0 (49.7)	60.5 (48.9)
Percent Black/Hispanic	11.0 (31.4)	6.7 (24.9)	8.2 (27.4)
Income (thousands)†	41.4 (11.3)	41.0 (10.3)	40.5 (10.1)
Self-reported health status**	1.94 (.87)	2.04 (.94)	2.06 (.91)
Worry about health††	2.56 (.84)	2.53 (.84)	2.52 (.84)

*Mean and (standard deviation)

†Mean household income in subject's telephone exchange

**Five-point scale; 1 = excellent, 5 = poor

††Four-point scale; 1 = not at all, 4 = most of the time

transformed probability estimates derived from the NHEFS logistic model. Logarithmic transformations were also applied to the CDC probabilities. CHD mortality rates increase monotonically with age and are considerably higher for men than women. To remove these effects from the validity assessments, partial correlations were also computed for each HRA score, adjusting for age, gender, and race.

Results

Baseline interviews were completed by 732 of 1,276 eligible adults (57.4 percent) contacted for the field trial. A logistic regression analysis of response rates indicated that the probability of obtaining a completed interview was not affected by demographic characteristics (gender, race, age, and average income level in the telephone exchange), self-reported health status, or concern about the state of one's health. Similarities between field trial participants and eligible subjects who declined to be interviewed are shown in Table 3. The median age of the sample was 38 years. Respondents were somewhat younger and healthier than the general population because persons with diabetes, hypertension, or evidence of overt cardiovascular disease were screened out of the sample.

Validity of CHD Risk Factors

The first set of analyses focused on five specific CHD risk items. These analyses were restricted to respondents whose physiological measurements were completed during the baseline interview. Table 4 shows the correlations between self-reported risk scores and validation measures for each of the four HRAs. These results reveal some sharp distinctions among the individual risk factors. The coefficients for cigarette smoking and relative weight were always .6 or greater for each of the four instruments. On the other hand, none of the correlations for physical activity, blood pressure, or cholesterol were higher than .5. The HRA risk scores for these items accounted for less than 25 percent of the variance in the physiological measures. The smallest correlations were found for blood pressure and cholesterol. This occurs primarily because many people do not know the appropriate values and either assume that their levels are average or make inaccurate guesses. The magnitudes of these correlations were relatively consistent from HRA to HRA; there was more variation among different risk factors than among instruments.

Validity of Heart Attack Risk Scores and Appraised Age

The second set of analyses examined the validity of HRA scores for heart attack risk and appraised age. There are two potential sources of error in total risk scores. First, respondents may be unaware of the appropriate values for physiological measures. When asked to specify exact values during in-person interviews, only 27 percent of the respondents gave systolic blood pressure values within 10 mmHg of those recorded by the field technicians, and only 4 percent estimated their total serum cholesterol to within 20 mg/dl of laboratory readings.

For self-scored appraisal instruments, a second source of error lies in computational mistakes made by respondents when calculating their total risk scores. Many of the field trial participants had difficulty computing these scores (Table 5). Error rates were affected both by the mathematical skills required by each instrument (addition, subtraction, or division) and by respondents' educational attainment (persons without four-year college degrees were twice as likely to make mistakes as those with baccalaureate or advanced degrees).

To assess the effects of these potential sources of error, separate correlations were computed for the risk scores reported by respondents, scores corrected for mathematical errors, and the scores that would have resulted from the physiological measurements taken by field technicians. The last of these three scores is similar to estimates based on physical examination data evaluated in our previous study.[6]

The upper panel of Table 6 shows the correlations for each HRA based on all respondents who were able to calculate risk scores and had a complete set of physiological measurements. The CDC risk appraisal had the highest correlations. This computerized instrument uses a more sophisticated scoring algorithm than the other HRAs and was the only one of the four that, like the logistic models, provides probability estimates for CHD mortality over a ten-year period. The CDC estimates, however, were consistently higher than the risks predicted by the NHEFS models (average risks of 1,522 per 100,000 compared with 1,173 per 100,000). Determine Your Medical Age had the next highest set of correlations, even though this instrument is not designed specifically for heart disease. Mathematical errors played an important role in the validity of this HRA, as the average coefficient increased from .41 to .55 when respondents' mistakes were corrected. Determine Your Medical Age also benefits considerably from the fact that CHD risk increases greatly with age, so that any HRA based on appraised age should have a reasonably high correlation in this type of analysis.

The pattern of results was considerably different for the Arizona Heart Institute HRA than for the other instruments. The correlations for this HRA were always about .52, regardless of whether computational errors were corrected or actual physiological values were substituted for the respondent's choices. This probably occurred because comparatively few people completing this instrument made errors in their risk calculations and because cholesterol values are not re-

TABLE 4—Correlations (and 95 Percent Confidence Intervals) between Validation Measures and Risk Item Scores

	Arizona Heart Institute (N = 67–68)		CDC (N = 63–77)		Determine Your Medical Age (N = 64–70)		RISKO (N = 55–64)	
Risk Items	r	95% CI	r	95% CI	r	95% CI	r	95% CI
Cigarettes Smoked per Day	.85	(.77–.91)	.66	(.52–.77)	.75	(.62–.84)	.80	(.69–.88)
Relative Weight	.74	(.61–.84)	.86	(.79–.91)	.64	(.47–.76)	.70	(.55–.81)
Physical Activity	–.27	(–.48 to –.04)	–.48	(–.64 to –.28)	–.47	(–.64 to –.26)	—	—
Systolic Blood Pressure	.47	(.26–.64)	.36	(.14–.54)	.00	(–.23–.24)	.50	(.29–.67)
Total Serum Cholesterol	—	—	.40	(.17–.59)	.18	(–.06–.41)	.26	(.00–.49)

TABLE 5—Percentage of Baseline Respondents Making Computational Errors When Calculating HRA Risk Scores

	Arizona Heart Institute (N = 157)	Determine Your Medical Age (N = 170)	RISKO (N = 155)
Score computed correctly	86.6	41.8	74.8
Score computed incorrectly	7.0	32.9	14.2
Could not compute score	6.4	25.3	11.0
Totals	100%	100%	100%

TABLE 6—Correlations between HRA Risk Scores and Natural Logarithm of NHEFS Ten-year CHD Mortality Risk Estimates

	Arizona Heart Institute (N = 124)	CDC (N = 137)	Determine Your Medical Age (N = 104)	RISKO (N = 112)
Zero-order correlations:				
Self-reported	.52	—	.41	.13
Corrected for errors	.52	.75	.55	.09
Physiologic field measures	.52	.77	.56	.40
R^2 for NHEFS estimate regressed on sex, age, and race	.85	.90	.87	.86
Partial correlations*:				
Self-reported	.42	—	.12	.19
Corrected for errors	.39	.47	.42	.39
Physiologic field measures	.39	.51	.43	.68

*Adjusted for respondent's gender, age, and race.

quired. RISKO had the smallest correlation with logistic estimates for the scores reported by respondents (.13), but in contrast to the other three HRAs, RISKO scores represent relative rather than absolute mortality risk.

A more refined assessment of validity can be made by examining partial correlations between HRA scores and the logistic estimates after controlling for age, gender, and race, as shown in the lower panel of Table 6. These three demographic characteristics consistently explained 85 to 90 percent of the variation in the log-transformed logistic rates. The resulting partial correlations were smaller than the unadjusted correlations in Table 5 for all HRAs except RISKO. The partials for self-reported Medical Age and RISKO scores were less than .2, but these coefficients more than doubled when mathematical errors were corrected. The magnitudes of the partial correlations also showed small improvements for three HRAs and a large increase for RISKO when physiological measurements taken in the field were substituted for the values selected by respondents.

Discussion

The results of this randomized field trial have several important implications for the accuracy of HRA instruments when used by a general population.

First, self-reported risk scores for cigarette smoking and relative weight appear to be accurate enough for use in HRAs. However, reports for physical activity, blood pressure, and cholesterol were frequently inaccurate when compared with more objective measures. Many people don't know their blood pressure or cholesterol levels and don't do a very good job of guessing what these values are. In the field trial, 27 percent of the sample reported systolic blood pressure values that were within 10 mmHg of the readings taken by field technicians and only 4 percent gave cholesterol values that were within 20 mg/dl.

The HRAs chosen for the field trial exhibited modest correlations with the CHD mortality risk estimates predicted by the NHEFS epidemiologic model. After adjusting for variations attributable to age, gender, and race, the partial correlations for all three of the self-scored instruments were approximately .4 when computational errors had been corrected. The partial correlation for the CDC HRA, the only instrument that produced probability estimates, was only slightly higher (.47). Moreover, like other HRAs based on the Geller-Gesner techniques, CDC's ten-year mortality risk predictions were consistently higher than those produced by the NHEFS data.[6,7] Several of the more recently developed HRAs incorporate logistic equations that may be expected to yield higher partial correlations than the field trial instruments.

The trial results also indicate that, when used in the field, the accuracy of the self-scored HRAs is attenuated by mathematical errors made by respondents and, to a lesser extent, by a lack of knowledge of the exact values for blood pressure, cholesterol, height, and weight. Computational errors substantially reduced the correlations for RISKO and Determine Your Medical Age, but not for the Arizona

Heart Institute instrument. This appears to be a function of the computational burden a self-scored HRA imposes on respondents. Determine Your Medical Age requires three mathematical operations (addition, subtraction, and division); RISKO, two (addition and subtraction), while the Arizona Heart Institute scores involve only addition. Computational errors and uncertainty regarding one's physiological status have also been found to lower the test-retest reliability of these HRAs.[8] Providing respondents with accurate physiological measurements and checking HRA calculations for errors will therefore improve both the reliability and the validity of self-scored risk estimates.

Finally, it should be noted that the analyses presented in this paper are restricted to CHD mortality. Since the risk factors for CHD are relatively well established, HRA risk estimates may be less accurate for other, less frequent causes of death and morbidity. ☐

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Clinical and Epidemiological Features of Neurotoxic Shellfish Poisoning in North Carolina

Peter D. Morris, MD, MPH, Douglas S. Campbell, MD, MPH, Ted J. Taylor, PhD, and John I. Freeman, DVM, MPH

ABSTRACT

Background: In October 1987, a red tide due to *P. brevis* affected the North Carolina coast for the first time. The purpose of our study was to describe the clinical and epidemiological features of neurotoxic shellfish poisoning (NSP), an illness caused by eating shellfish contaminated with the neurotoxins of *P. brevis.*

Methods: Active surveillance was established for cases of NSP. A descriptive study of the NSP cases was then completed.

Results: Forty-eight persons, who had eaten oysters at 20 meals, met the case definition. A variety of gastrointestinal tract and neurological symptoms were reported. The illnesses were generally mild and of short duration, and there were no deaths. Forty-one (85 percent) affected persons lived in five communities located within a 70-kilometer area along the coast. Cases occurred from October 27 to December 9; 27 (56 percent) of the cases occurred before the first closure of affected shellfish waters on November 2. There was a significant increase in the illness attack rate with an increase in the number of oysters eaten.

Conclusions: Routine monitoring of coastal waters for *P. brevis* is needed to facilitate earlier recognition of red tides, closure of affected areas, and education of the public before substantial exposure to contaminated shellfish occurs. (*Am J Public Health* 1991;81:471–474)

Introduction

Neurotoxic shellfish poisoning (NSP) is an unusual syndrome characterized by paresthesia, reversal of hot and cold temperature sensation, myalgia, vertigo, and ataxia.[1–5] Other symptoms include abdominal pain, nausea, diarrhea, burning pain in the rectum, headache, bradycardia, and dilated pupils.[1–5] It is caused by eating filter-feeding shellfish (oysters, clams, coquinas, and other bivalve mollusks)[5] containing neurotoxins produced by the marine dinoflagellate, *Ptychodiscus brevis*[6] (formerly *Gymnodinium breve*[7]). Accumulation of neurotoxins in fish probably does not occur.[5]

P. brevis can also cause a respiratory irritation syndrome.[8] The dinoflagellates are easily lysed by the surf, releasing toxins that cause conjunctival irritation, rhinorrhea, and nonproductive cough in persons who are close to the water.

The health problems caused by *P. brevis* are associated with a red tide bloom.[3,4,8] Red tide is a discoloration of sea water caused by a proliferation of large numbers of dinoflagellates.[8] In late October 1987, the North Carolina coast was affected by a *P. brevis* red tide, the first such episode outside of Florida and Texas.[9] The source of the dinoflagellates was probably a red tide that occurred off the southwest coast of Florida in September. The organism was carried around Florida by the Florida current and then north along the east coast of Florida by the Gulf Stream. In late October, a strong inshore eddy of the Gulf Stream, later detected using satellite imagery, probably brought algae-laden water to the North Carolina coast. Under favorable reproductive conditions, the organism proliferated, resulting in an extensive red tide. By October 30, the characteristic discoloration, massive fish kills, and cases of the respiratory irritation syndrome were reported along the North Carolina coast. On November 2, the organism was identified, and the affected shellfish areas were closed. On November 4, the first official news release warned the public not to consume shellfish from affected areas.

In early November, the Environmental Epidemiology Section of the North Carolina Department of Environment, Health, and Natural Resources learned of several neurological illnesses that occurred following consumption of shellfish harvested from known red-tide-affected waters. To detect other possible cases of NSP in North Carolina, active surveillance was established. In this report, we describe the general clinical and epidemiological features of NSP cases detected in this manner.

Methods

Telephone contact was made with coastal county primary-care physicians, health departments, and emergency rooms; emergency rooms in referral hospitals for the coastal counties; and the

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Address reprint requests to Peter D. Morris, MD, MPH, Environmental Epidemiology Section, Division of Epidemiology, North Carolina Department of Environment, Health, and Natural Resources, P.O. Box 27687, Raleigh, NC 27611–7687. Dr. Morris, with the CDC Epidemiology Program Office, Division of Field Services, Atlanta, was on assignment to the State of North Carolina at the time of this study. Co-authors Campbell, Taylor, and Freeman are all with the NC Department of Environment, Health, and Natural Resources. This paper, submitted to the Journal December 15, 1989, was revised and accepted for publication August 13, 1990.

TABLE 1—Symptoms of 48 Neurotoxic Shellfish Poisoning Cases, North Carolina, 1987

Symptom	N (%) of Patients with Symptom	Median Time between Eating Shellfish and Onset Hr. (Range)	Median Duration Hr. (Range)
Gastrointestinal			
Abdominal pain	23 (48)	5 (1–18)	17 (1–48)
Nausea	21 (44)	3 (0.5–18)	18 (1–48)
Diarrhea	16 (33)	10 (1–18)	17 (2–48)
Vomiting	5 (10)	3 (1–15)	8 (0.5–48)
Neurological: sensation			
Paresthesia	39 (81)	4 (0.25–14)	8 (0.5–51)
Temperature reversal	8 (17)	7 (0.25–16)	8 (1–26)
Myalgia	6 (13)	6 (0.5–13)	7 (2–24)
Neurological: cerebellum			
Vertigo	29 (60)	3 (0.25–17)	16 (0.5–72)
Ataxia	13 (27)	3 (0.25–14)	7 (0.5–40)
Tremor	3 (6)	3 (0.25–10)	5 (0.5–9)
Neurological: motor			
Weakness	15 (31)	3 (0.25–17)	24 (0.5–51)
Systemic			
Malaise	24 (50)	3 (0.25–15)	24 (2–72)
Chills	10 (21)	2 (2–8)	11 (2–48)
Headache	7 (15)	2 (0.5–12)	7 (2–24)

North Carolina Division of Marine Fisheries. Each facility was asked to report any neurological or gastrointestinal (GI) illnesses associated with the consumption of shellfish after October 23 (when the toxic dinoflagellate probably first arrived on shore). Telephone interviews were completed on persons reported through surveillance and persons identified by the surveillance cases who had eaten at the same shellfish meals. The following data were sought from each person: name; age; sex; race; weight; presence of a chronic illness; medication use; date, time, and location of harvest or purchase of shellfish; type of preparation of shellfish; date and time shellfish were eaten; type and amount of shellfish consumed; time of onset of symptoms; estimates of duration of symptoms; estimates of days of bed rest and days lost from work or school; when and where medical attention was sought; and type of medical treatment. The questionnaire included a list of symptoms previously associated with NSP.[2] In addition, the questionnaire included symptoms associated with two similar toxic dinoflagellate syndromes—paralytic shellfish poisoning and ciguatera fish poisoning.[2]

Shellfish (oysters, clams, or mussels) were considered toxic if they produced at least one case of illness characterized by any of the previously reported neurological symptoms of NSP—paresthesia, temperature sensation reversal, myalgia, vertigo, or ataxia. In addition, shellfish must have been harvested from a known red-tide-affected area (>5000 *P. brevis* organisms/liter of water detected by direct microscopic examination). A person who had eaten toxic shellfish was defined as a case if any of the previously reported neurological symptoms of NSP occurred within 24 hours of eating the shellfish and persisted for at least one hour.

A retrospective cohort analysis was completed on those persons who had eaten at toxic shellfish (oyster) meals. Fisher's exact test was used to compare illness attack rates in those exposed and unexposed to toxic oysters.[10] Among persons who had eaten toxic oysters, summary 95 percent confidence intervals were used to assess the effect of potential risk factors (number of shellfish eaten, age, sex, presence of a chronic illness, chronic use of a medication, or alcohol consumption with the meal) on the attack rate of NSP.[11] The Mantel-Haenszel chi-square test for trend was used to assess the change in attack rates with an increase in the number of shellfish eaten.[12]

Residual specimens of toxic oysters were collected whenever available for laboratory analysis by the Environmental Sciences Section, Division of Laboratory Services, of the Department. Shellfish specimens harvested from the same general areas as those that were eaten at all the other shellfish meals were also collected for analysis. Since a bioassay specific for *P. brevis* neurotoxins was not available, a standardized mouse bioassay was used on the specimens to test for the presence of neurotoxicity.[13] This bioassay is based on the time until death of mice injected intraperitoneally with crude toxin residues extracted from shellfish. Relative toxicity is expressed in mouse units (MU). One mouse unit is the amount of crude toxin residue that, on average, will kill 50 percent of test mice in 930 minutes. Any detectable level of toxin per 100 grams of shellfish meat was considered a positive test.

Results

Forty-eight persons, who had eaten at 20 different meals, met the case definition. Two other persons experienced several of the typical NSP neurological symptoms after eating oysters in two different restaurants, but they were excluded from analysis because shellfish harvest sites could not be determined.

The median incubation period was three hours (range, 15 minutes to 18 hours). GI and neurological symptoms tended to occur at the same time. The constellation of symptoms presented by the cases is shown in Table 1. Forty-five (94 percent) had multiple symptoms, and 34 (71 percent) had more than one neurological symptom. There were no instances of paralysis or respiratory arrest. However, a 43-year-old White female was admitted to an intensive care unit for severe symptoms of bilateral carpopedal tremor and myalgia; total body paresthesia; ataxia; and vertigo. These symptoms developed 9.5 hours after eating approximately 45 oysters. On neurologic examination, her physician found decreased sensation, decreased reflexes, and temperature reversal in both lower extremities. She was treated supportively and recovered completely, approximately nine hours after onset of illness.

The median duration of illness was 17 hours (range, one to 72 hours). Twelve affected persons (25 percent) required one day of bed rest, while only one required two days of bed rest. Six (12.5 percent) missed one day of either work or school. One patient was hospitalized, but there were no deaths.

The 48 affected persons had a median age of 43 years (range, nine to 66 years); 27 (56 percent) were male, and all were white. Forty-one of the affected persons (85 percent) lived in five communities located along a 70-kilometer coastal area in Carteret and Onslow counties. This was the first area to be affected by the red tide, and it consistently had some of the highest *P. brevis* cell concentrations throughout the bloom. Cases occurred from October 27 to December 9 and peaked on Novem-

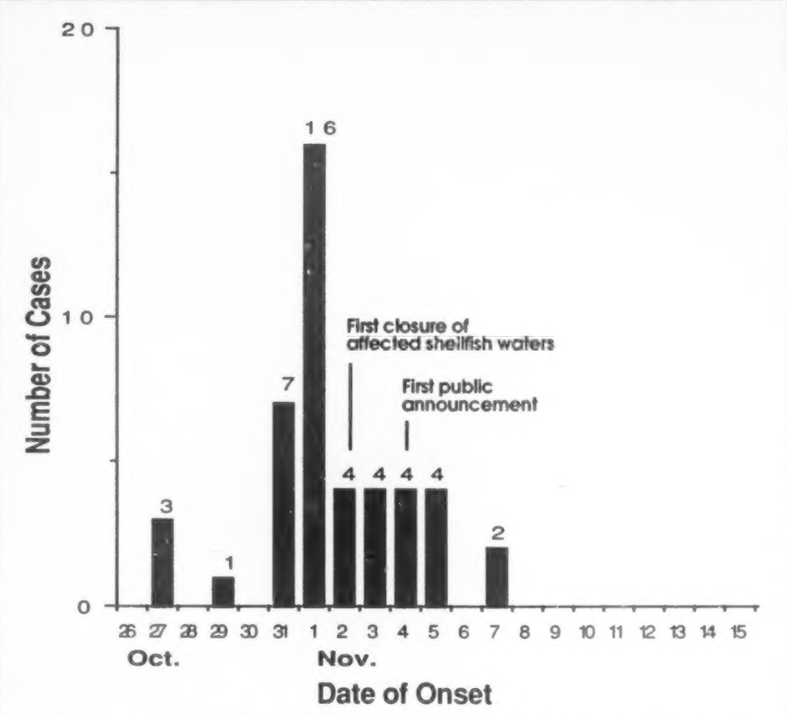

FIGURE 1—Date of Illness Onset, for 48 Neurotoxic Shellfish Poisoning Cases, North Carolina, 1987.

ber 1 (Figure 1). Three additional persons (not shown in figure) had eaten oysters harvested on October 31 and kept frozen until the date of the meal, December 8, when they became ill. Twenty-seven of the cases (56 percent) occurred before the first shellfish water closure on November 2.

All of the shellfish eaten at the meals were harvested within the same 70-kilometer stretch of coast discussed above. Forty-four of the affected persons (92 percent) had eaten shellfish harvested either by themselves or by another individual (relative, friend, or neighbor); whereas four (8 percent) had purchased oysters from fish markets. Two of the four had purchased oysters from a market before the shellfish harvest ban and the first public news release; the remaining two had purchased oysters from a fish market shortly after the ban and the news release.

Oysters were the only shellfish eaten at 19 of the 20 meals, while both oysters and clams were eaten at one meal. All cases were associated with freshly harvested shellfish, except for the one instance in December where frozen oysters were consumed. The shellfish consumed were generally described as being normal in appearance, taste, and smell. In all but one instance, the shellfish had been cooked before eating; cooking methods included steaming, roasting, baking, boiling, and frying.

A total of 98 persons had attended the 20 shellfish meals. The retrospective cohort analysis of these 98 persons revealed an illness attack rate of 56 percent (48/85) among those who had eaten oysters; none of the 13 persons who had not eaten shellfish became ill. None of the 25 case household members, who were not present at the shellfish meals, developed neurological or GI illnesses suggestive of NSP. Among those who had eaten oysters, the number of shellfish eaten was the only potential risk factor examined that significantly influenced the attack rate of NSP. The risk of developing NSP increased with the number of oysters eaten; this trend in disease risk was statistically significant (Table 2). The attack rates and relative risks in Table 2 were essentially unchanged after adjusting for age and the other potential risk factors examined in this study. Thus, only unadjusted data are shown in Table 2.

Samples of oysters eaten by four affected persons from two meals tested positive for *P. brevis* neurotoxins (35 and 60 MU, respectively). Samples of oysters harvested from the same general areas as those that were eaten by all the other cases also tested positive (mean, 62 MU; range, 48–170 MU).

Discussion

Since it was not possible to confirm the NSP cases with a specific laboratory test, some of the reported illnesses may have been due to other causes. However, this possibility was minimized by using a specific case definition—illness typical of previously reported NSP after eating shellfish harvested from a known red tide-affected area. In addition, epidemiological findings support the conclusion that the illnesses reported were due to NSP. First, none of the reporting sources were aware of any increase in neurological or GI illnesses unrelated to shellfish consumption during the time of the red tide bloom. Second, none of the 13 persons who had not eaten shellfish at any of the 20 meals developed neurological or GI symptoms suggestive of NSP. Third, none of the 25 case household members who were not present at the shellfish meals developed illnesses suggestive of NSP. Finally, among those who had eaten oysters, the illness attack rate rose with the number of oysters eaten. In addition to this epidemiological evidence, sample oysters tested positive for *P. brevis* neurotoxins, using the mouse bioassay.[13]

The affected persons developed a variety of neurological and GI symptoms similar to NSP cases previously reported from Florida.[3,4] However, we did find differences in the North Carolina cases that could provide insight in future clinical evaluations. Nausea and vertigo were more common in North Carolina than in Florida. In addition, some symptoms reported in North Carolina were not mentioned in the Florida reports; these included malaise, muscle weakness, tremor, and a choking sensation or throat tightness (this symptom was not included on the questionnaire but was spontaneously reported by several persons). Finally, 21 of the North Carolina cases (44 percent) had latency periods greater than three hours; whereas in Florida, all NSP cases had latency periods of less than three hours. These longer latency periods may have been due to differences in the relative concentrations of *P. brevis* neurotoxins. Indeed, analysis of neurotoxins from water samples taken during the North Carolina red tide bloom demonstrated a greater proportion of a slower acting component than had been found in a previous Florida red tide.[14]

TABLE 2—Number of Oysters and Risk of Neurotoxic Shellfish Poisoning, for Those Who Consumed Oysters, North Carolina, 1987

Number of Oysters	Cases	Total	AR (%)*	RR	95% CI
<12	2	15	13.3	1.0	
12–35	16	30	53.3	4.0	1.1–15.2
≥ 36	29	39	74.4	5.6	1.5–20.5

*AR = Attack Rate
Mantel-Haenszel chi-square = 15.9 $p < 0.000034$

Even though *P. brevis* was detected in shellfish waters until mid-February 1988, the number of cases dropped sharply on November 2, suggesting that the shellfish water closure and public news release were effective in preventing illness. Most of the cases occurred among persons who had eaten oysters harvested either by themselves or by a relative or friend, emphasizing the need to educate not only seafood merchants but also the general public about the dangers of harvesting and consuming shellfish from red tide-affected waters. Furthermore, most of the NSP cases lived along that section of the coast where the bloom started and where *P. brevis* reached its highest concentrations. These persons had easier access to toxic oysters, which is possibly the cause of this geographic clustering. It is therefore important to focus education measures on populations of coastal areas directly affected by red tide waters.

The effectiveness of shellfish water closures and public education as preventive measures depends on early detection of *P. brevis* in water samples. In North Carolina, more than a week passed from the time dinoflagellate cells were first carried to the coast until the first shellfish water closure and public news release. The characteristic discoloration of water, fish kills, and the respiratory irritation syndrome indicate the presence of a red tide and possible shellfish contamination and are thus useful warning signs. However, Florida officials, by routinely sampling coastal waters during the months when red tides tend to be more common, can usually detect levels of the organism high enough to contaminate shellfish (>5,000 organisms/liter of water), but not yet high enough to cause noticeable effects in the water (Florida Department of Natural Resources contingency plan for control of shellfish potentially contaminated by marine biotoxins. Bureau of Marine Research, St. Petersburg, Florida, 1985). This probably prevents cases of NSP.

Routine monitoring of coastal waters has now been established in North Carolina. The Gulf Stream is checked for inshore incursions or eddies by examining oceanographic frontal analyses received from the National Hurricane Center in Miami. These analyses are prepared thrice weekly from hourly infrared images taken by the Geostationary Operational Environmental Satellite series. These analyses are normally used to assist mariners who wish to either utilize or avoid the swiftly flowing Gulf Stream. Fishermen also use the analyses to pinpoint potentially good fishing areas. If a strong eddy occurs along the North Carolina coast, then sampling of relevant coastal waters for *P. brevis* is started immediately. Such a monitoring program should facilitate earlier recognition of possible future red tides, closure of affected areas, and education of the public before substantial exposure to contaminated shellfish occurs. ☐

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Bone Cancer Incidence Rates in New York State: Time Trends and Fluoridated Drinking Water

37

ABSTRACT

Background: Recent animal studies of the potential carcinogenicity of fluoride prompted an examination of bone cancer incidence rates.

Methods: Trends in the incidence of primary bone cancers, including the incidence of osteosarcomas were examined among residents of New York State, exclusive of New York City. Average annual osteosarcoma incidence rates in fluoridated and non-fluoridated areas were also compared.

Results: Among persons less than 30 years of age at diagnosis, bone cancer incidence among males demonstrated a significant increase since 1955, while incidence among females has remained unchanged. A significant decrease in bone cancer incidence rates since 1955 was observed among both males and females age 30 years and over at time of diagnosis. Osteosarcoma incidence rates have remained essentially unchanged since 1970, among both younger and older males and females. The average annual age adjusted incidence of osteosarcomas (1976–1987) in areas served by fluoridated water supplies was not found to differ from osteosarcoma incidence rates in non-fluoridated areas.

Conclusions: These data do not support an association between fluoride in drinking water and the occurrence of cancer of the bone. (*Am J Public Health* 1991;81:475–479)

Martin C. Mahoney, PhD, Philip C. Nasca, PhD, William S. Burnett, MD, DrPH, and James M. Melius, MD, DrPH

Introduction

Artificial fluoridation of public drinking supplies has resulted in an ongoing debate concerning the potential health effects of this practice. Proponents have emphasized the benefits to dental health from the artificial fluoridation of drinking water, while opponents have suggested that a variety of adverse health outcomes, including cancer, can result from this exposure.

Interest in the potential carcinogenicity of fluoride has resurfaced as a result of recently released findings from a National Toxicology Program (NTP) study of rodents with life-long exposures to fluoride. That study reported an increased incidence of osteosarcomas among male rats ingesting a high fluoride diet. Although the significance of these findings is unclear, the study has raised renewed questions about the association between fluoridated drinking water and bone cancers, particularly osteosarcomas, in human populations.

Several studies which have compared cancer mortality in fluoridated and non-fluoridated areas have determined that cancer mortality is unrelated to fluoride levels in drinking water.[1–6] However, the relationship between cancer incidence and fluoridated water has not been thoroughly examined. We were able to identify only two previous studies of cancer incidence and fluoridated water supplies.[1,7] Using data from the Second and Third National Cancer Surveys, Hoover *et al*,[1] compared cancer incidence for a fluoridated (Denver) and nonfluoridated metropolitan area (Birmingham). No significant differences in overall cancer incidence were apparent, nor were significant differences in bone cancer incidence observed. Kinlen and colleagues[7] reported no differences in the incidence of bone cancers among areas in England and Wales, and the United States, Holland, and New Zealand, with high and low fluoride levels in drinking water. An additional study of exposure to fluoride among an occupational cohort reported an elevated cancer incidence among exposed workers, attributable to an excess of respiratory cancers.[8] It is not possible to rule out smoking as the cause of this respiratory cancer excess since information on smoking histories was unavailable. No cases of bone cancer were reported in this study of occupational fluoride exposure.

Widespread artificial fluoridation of drinking water supplies in New York State began primarily during the 1950s and early 1960s. By analyzing data in the New York State Cancer Registry, it is possible to compare current bone cancer incidence rates with rates prior to widespread fluoridation. This paper presents data on time trends in bone cancer incidence, including the incidence of osteosarcomas, among residents of New York State, exclusive of New York City, and compares average annual osteosarcoma incidence rates in fluoridated and nonfluoridated areas.

Methods

The New York State Cancer Registry has been operational since 1940, when

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From the Bureau of Cancer Epidemiology and the Division of Occupational Health and Environmental Epidemiology, New York State Department of Health, and the School of Public Health at State University of New York at Albany. Address reprint requests to Martin C. Mahoney, PhD, Research Scientist, Bureau of Cancer Epidemiology, New York State Department of Health, Empire State Plaza, Tower 565, Albany, NY 12237-0683. This paper, submitted to the Journal June 25, 1990, was revised and accepted for publication October 18, 1990.

cancer became a reportable disease in NYS. Beginning in 1973, this population-based registry expanded coverage to include the New York City area, thereby encompassing all of NYS. While physicians, laboratories, and hospitals are required to report cancer cases, most cases in the NYS Cancer Registry and other cancer registries are reported by hospital tumor registrars and medical record administrators. Based on the results of special studies of hospital records, it is estimated that over 95 percent of all cancer cases in New York State, exclusive of New York City, are reported to the registry.[9]

Analyses were restricted to primary malignancies of the bone occurring among residents of New York State, exclusive of New York City, between 1955 and 1987. The New York City area was excluded due to lack of cancer data before 1973. Bone malignancies were coded to International Classification of Diseases, Ninth Revision (ICD-9),[10] site 170. Site 170 includes cancers of the bone and articular cartilage, but excludes carcinomas of any type other than intraosseous or odontogenic. Three separate comparisons of bone cancer incidence were undertaken.

The first comparison involved the calculation of average annual age-specific bone cancer incidence rates, between 1950 and 1987, for both males and females. Because of the distinct bimodal age pattern for bone cancers,[11] incidence rates were compared within two broad age categories: less than 30 years and 30 years and older at time of diagnosis. United States Census data were used to provide decennial (1960, 1970, 1980) population counts by age and sex for NYS,[12–14] while age- and sex-specific population estimates for intercensual years were furnished by the New York State Department of Health Bureau of Biometrics. Trends in incidence rates were evaluated by considering the slope of the regression line (t-test) using Statistical Analysis System (SAS) computer software.[15] Ninety-five percent confidence intervals for incidence rate ratios were calculated as described by Rothman.[16]

A second comparison involved the calculation of average annual age-specific osteosarcoma incidence rates, among males and females between 1970 and 1987. Histologic codes were assigned according to the Manual of Tumor Nomenclature and Coding[17] or more recently the International Classification of Diseases for Oncology.[18] Due to a different histologic classification scheme in use before 1970 it was not possible to distinguish histologic subtypes of bone cancers reported to the Registry prior to 1970. Trends in osteosarcoma incidence rates were also evaluated using the procedures described above.

Drinking water supplies in New York State contain no appreciable levels of naturally occurring fluoride. Thus, fluoridated water supplies are all artificially fluoridated. A third comparison evaluated average annual age-specific incidence rates for bone cancers and osteosarcomas, between 1976 and 1987, among males and females residing in fluoridated and nonfluoridated communities. Two metropolitan areas in New York State have maintained nonfluoridated drinking supplies: the City of Albany and Suffolk County.[19] Since less than 10 percent of the Nassau County population consumes fluoridated water, this metropolitan area was also considered to be nonfluoridated for the purposes of this study. Together, the 1980 population of these nonfluoridated areas (Albany, Suffolk County, Nassau County) was 2,707,540.[13] Since the nonfluoridated areas are urbanized communities, osteosarcoma incidence rates in these areas were compared with incidence in two groups of fluoridated communities: 1) counties located within standard metropolitan statistical areas (i.e. urbanized areas), and 2) counties not located within standard metropolitan statistical areas. Residents of nonmetropolitan areas are less likely to be served by a public water supply. Since these incidence rates were based on a limited number of cases (due to the infrequent occurrence of osteosarcomas), the standard error (SE) of each incidence rate was estimated by dividing the rate by the square root of the number of cases upon which it is based.[20]

Results

Between 1955 and 1987, 3,166 bone cancers were diagnosed among residents of New York State, exclusive of New York City (1,824 males, 1,292 females). About one-third (n = 1,100) of these cases occurred among persons less than 30 years of age at diagnosis (652 males, 448 females); 2,016 cases occurred among persons ages 30 and over at time of diagnosis (males 1,172, females 844).

Trends in Bone Cancer Incidence Rates

For bone cancers among persons less than 30 years of age at diagnosis, the slope of the regression line was significant among males ($p = 0.002$) with an observed increase in bone cancer incidence of 54 percent during the 1955–87 period. Among females (<30 years), bone cancer incidence increased 44 percent; however, the slope of the regression was not significant at the 5 percent level ($p = 0.07$). Further exploration of trends in bone cancer incidence among persons less than 30 years of age at time of diagnosis reveals that, relative to the period 1955–63, a significant increase in bone cancer incidence among males occurred only after 1969, while among females incidence was not significantly increased until the 1983–87 period (Table 1).

Bone cancer incidence rates among persons age 30 years and over at diagnosis show a significant decrease among both males (44 percent decrease, $p = 0.0001$) and females (33 percent decrease, $p = 0.0001$). Relative to incidence during the period 1955–1963, bone cancer incidence rates among males were significantly decreased during the period 1970–75 and thereafter (Table 1). Among females age 30 years and older at diagnosis, significant decreases in bone cancer incidence rates did not occur until after 1975.

Trends in Osteosarcoma Incidence Rates

Among persons less than 30 years of age at time of diagnosis, 196 and 113 osteosarcomas were diagnosed among males and females, respectively, between 1970 and 1987. While there was some variation in annual osteosarcoma incidence rates, between 1970 and 1987, among persons less than 30 years of age at diagnosis, the slope of the regression line was not significant among either males (16 percent increase, $p = 0.71$) or females (18 percent increase, $p = 0.45$). As shown in Table 2, osteosarcoma incidence rates did not differ significantly during three six-year periods between 1970 and 1987.

Between 1970 and 1987, 135 osteosarcomas were diagnosed among males and 103 osteosarcomas were diagnosed among females age 30 years or older at time of diagnosis. Based on the slope of the regression line, osteosarcoma incidence rates have been stable among both males (16 percent decrease, $p = 0.58$) and females (24 percent increase, $p = 0.29$) over 30 years of age at diagnosis. Osteosarcoma incidence rates did not differ significantly during three six-year periods between 1970 and 1987 (Table 2).

TABLE 1—Rate Ratios by Sex and Age at Diagnosis for Bone Cancer Incidence Rates per Million among Residents of New York State, Exclusive of New York City, 1955–87

	<30 Years at Diagnosis							
	Males				Females			
Year of Diagnosis	No.	Rate	RR*	(95% CI)**	No.	Rate	RR	(95% CI)
1955–1963	116	6.1	1.00		91	4.8	1.00	
1964–1969	111	7.3	1.20	(0.93, 1.56)	86	5.7	1.19	(0.88,1.60)
1970–1975	132	8.2	1.34	(1.04, 1.72)	88	5.5	1.15	(0.86,1.54)
1976–1981	156	10.0	1.64	(1.29, 2.10)	85	5.6	1.17	(0.87,1.57)
1982–1987	137	9.4	1.54	(1.22, 1.97)	98	6.9	1.44	(1.09,1.94)
t-test for trend:***	p = 0.002						p = 0.07	

	30+ Years at Diagnosis							
	Males				Females			
Year of Diagnosis	No.	Rate	RR	(95% CI)	No.	Rate	RR	(95% CI)
1955–1963	392	19.8	1.00		245	11.4	1.00	
1964–1969	253	18.2	0.92	(0.78, 1.08)	181	11.7	1.03	(0.84,1.24)
1970–1975	204	14.3	0.72	(0.61, 0.86)	154	9.4	0.82	(0.68,1.01)
1976–1981	150	10.1	0.51	(0.42, 0.61)	125	7.2	0.63	(0.51,0.79)
1982–1987	173	11.0	0.56	(0.46, 0.66)	139	7.6	0.67	(0.54,0.82)
t-test for trend:	p = 0.0001						p = 0.0001	

*RR, rate ratio relative to 1955–1963 period.
**95% confidence interval.
***t-test for trend using annual incidence rates between 1955 and 1987.

TABLE 2—Rate Ratios by Sex for Osteosarcoma Incidence Rates per Million among Residents of New York State, Exclusive of New York City, 1970–87

	<30 Years at Diagnosis							
	Males				Females			
Year of Diagnosis	No.	Rate	RR*	(95% CI)**	No.	Rate	RR	(95% CI)
1970–1975	61	3.8	1.00		35	2.2	1.00	
1976–1981	71	4.6	1.21	(0.86,1.70)	41	2.7	1.23	(0.78,1.93)
1982–1987	64	4.4	1.16	(0.84,1.64)	37	2.6	1.18	(0.77,1.87)
t-test for trend:***	p = 0.71						p = 0.45	

	30+ Years at Diagnosis							
	Males				Females			
Year of Diagnosis	No.	Rate	RR	(95% CI)	No.	Rate	RR	(95% CI)
1970–1975	45	3.2	1.00		28	1.7	1.00	
1976–1981	48	3.2	1.00	(0.68,1.54)	36	2.1	1.24	(0.75,1.98)
1982–1987	42	2.7	0.84	(0.56,1.28)	39	2.1	1.24	(0.79,1.95)
t-test for trend:	p = 0.58						p = 0.29	

*RR, rate ratio relative to 1970–1975 period.
**95% confidence interval.
***t-test for trend using annual incidence rates between 1970 and 1987.

Bone Cancer, Osteosarcoma Incidence and Fluoridated Drinking Water

As shown in Table 3, average annual sex- and age-specific bone cancer and osteosarcoma incidence rates (1976–87) did not differ between areas with and without fluoridated drinking water supplies.

Discussion

Although changes in bone cancer incidence since 1955 are evident, these changes have not been consistent between age groups or between the two genders. Bone cancer incidence rates among males less than 30 years at diagnosis have increased since 1955, while no statistically significant change in incidence was observed among females in this age group. Among persons ages 30 years and older, significant decreases in bone cancer incidence were noted among both males and females since 1955. Incidence increases, which have occurred, have been re-

TABLE 3—Age- and Sex-specific Incidence Rates per Million for Bone Cancers and Osteosarcomas among Residents of New York State, Exclusive of New York City, 1975–87, by Age at Diagnosis

	Bone Cancer Incidence								
	Nonfluoridated Areas			Fluoridated SMSA* Counties			Fluoridated NonSMSA Counties		
Age at Diagnosis	No.	Rate	(SE)**	No.	Rate	(SE)	No.	Rate	(SE)
<30 years									
males	75	9.7	(1.12)	158	9.0	(0.71)	60	11.3	(1.46)
females	51	6.7	(0.94)	110	6.4	(0.61)	22	4.3	(0.92)
30+ years									
males	98	12.4	(1.25)	181	10.4	(0.78)	44	9.0	(1.36)
females	64	7.0	(0.87)	162	8.0	(0.62)	38	6.7	(1.12)

	Osteosarcoma Incidence								
	Nonfluoridated Areas			Fluoridated SMSA* Counties			Fluoridated NonSMSA Counties		
Age at Diagnosis	No.	Rate	(SE)**	No.	Rate	(SE)	No.	Rate	(SE)
<30 years									
males	34	4.4	(0.75)	75	4.3	(0.43)	26	4.9	(0.96)
females	24	3.2	(0.64)	44	2.5	(0.38)	10	2.0	(0.62)
30+ years									
males	26	3.3	(0.64)	50	2.9	(0.41)	14	2.9	(0.77)
females	22	2.4	(0.51)	45	2.2	(0.33)	8	1.5	(0.51)

*SMSA = Standard Metropolitan Statistical Areas.
**SE = standard error.

stricted to more recent time periods among the cohort of persons born after the initiation of widespread artificial fluoridation of water supplies suggesting that any potential association with fluoride must be differential by age. The increased bone cancer risk observed in younger persons could be consistent with lifelong exposures to fluoride including exposures at times when growth rates for skeletal tissue are at a peak. However, other findings from this study do not support an association between exposure to fluoridated water and bone cancers. The incidence of bone cancers and osteosarcomas was not found to differ between areas with and without fluoridated drinking water supplies.

The potential specificity of effect from fluoride exposure was also examined by evaluating trends in the incidence of osteosarcomas. The incidence of osteosarcomas reveal no significant changes among either males or females, in either age group, since 1970. An examination of osteosarcoma incidence, between 1976 and 1987, in areas which have remained nonfluoridated did not indicate any significant differences compared to areas supplied by fluoridated water sources.

The limited secular increases observed in the number of bone cancer cases and osteosarcoma cases suggest that any potential effect of fluoride is probably minimal. Among population subgroups demonstrating incidence increases, the estimated annual increase in rates of bone cancers and osteosarcomas was 1–2 percent. Despite the ubiquitous nature of fluoride exposures, bone cancer and osteosarcomas remain as relatively infrequent cancers.

As with most previous studies of potential health effects resulting from exposure to fluoridated drinking water, the present study utilizes an ecological design and lacks precise measures of exposures to fluoridated drinking water, and other fluoride sources, at an individual level. Further, this study has not attempted to account for population mobility and movement between fluoridated and nonfluoridated areas. While most fluoridation of water supplies occurred during the 1950s and 1960s, some communities across New York State have initiated fluoridation of drinking water supplies at later times.

Although these data do not demonstrate consistent shifts in the incidence of bone cancers or osteosarcomas, recent increases in bone cancer incidence relative to the 1955–63 period are evident, particularly among younger persons. At this time it is not possible to establish an association between fluoride and bone cancer incidence. The potential role of fluoride in the incidence of bone cancers might be more thoroughly investigated using an analytic study design allowing more careful assessment of residential history to minimize misclassification of exposure to fluoridated drinking water and an examination of fluoride exposure from sources other than drinking water. ☐

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Federal Hazardous Waste Law Plagued by Ambiguities, says Report

A portion of the 1986 federal law requiring each of the 50 states to ensure the availability of hazardous waste treatment or disposal facilities over the next 20 years has been plagued by ambiguities, unanswered policy questions, and at least one lawsuit, according to a report to the US Environmental Protection Agency (EPA) by a Tufts University researcher. The report by Sharon N. Green, senior environmental research analyst at Tufts' Center for Environmental Management (CEM), recommends that Congress reassess the need for hazardous waste treatment and disposal facilities, given uncertainties about waste generation and management over the next 20 years.

The report cites data from 13 Northeast states projecting an 18 percent decline in hazardous waste generation by the year 2009, based on economic forecasting, waste reduction programs, expected new regulations, and contaminated site cleanups.

In her report, Green suggests a 15-point set of recommendations, including that the goals of the current provision be clarified to explicitly address waste reduction, facility siting, and attaining geographic equity for waste management services.

Other recommendations in the report, released in December 1990, include:

- Congress should consider state and federal roles in providing hazardous waste management services, given that responsibility for waste generation and management currently rests with the private sector.
- Congress should reconsider the appropriateness of threatening non-complying states with the loss of federal funds for hazardous waste cleanup projects.
- EPA should provide additional guidance to states in preparing hazardous waste projections.
- Data collection and planning methods among the states should be standardized.

"Capacity assurance planning probably will become a viable public policy only if significant revisions are made to the structure and function of the provision and implementation process," Green concludes in the report.

Public Health Briefs

Neonatal Diarrhea at a Maternity Hospital in Rangoon

Daw Tin Aye, MBBS, Dip Bact, David A. Sack, MD, I. Kaye Wachsmuth, PhD, Daw Thein Kyi, MBBS, DCH, and Shin Min Thwe, MBBS

ABSTRACT

Between 1981 and 1986, 1,540 infants born at the Central Women's Hospital in Rangoon were transferred to the Sick Baby Unit because of diarrhea (15.4 per 1,000 live births). Rates among cesarean infants were five times as high as those of infants born vaginally (51.0 and 10.3 per 1000 live births, respectively). One hundred eighty-four of the infants with diarrhea died (case fatality rate = 12 percent). We conclude that neonatal diarrhea is endemic in this large maternity hospital in Burma, and that control efforts should be targeted especially to cesarean and low birthweight infants. (*Am J Public Health* 1991; 81:480–481)

Introduction

Neonatal diarrhea was an important problem in American and Western European hospitals in the 1940s and 50s[1-6]; in recent years it has become relatively rare. The reported episodes of diarrhea tended to occur in clusters and were often associated with particular serotypes of *E. coli* which became known as enteropathogenic *E. coli* (EPEC).[7-10]

In developing countries diarrhea is a major endemic public health problem with an estimated five million deaths annually.[11] The bulk of the cases occur between the ages of six months and five years and most strategies for diarrheal disease control are targeted to this age group. Relative to this massive public health problem, neonatal diarrhea occurring in children born in hospital appears to be a relatively minor one.

Nevertheless, increasing numbers of children are born in developing country hospitals. A previous three-month survey in two Rangoon maternity hospitals found neonatal diarrhea to be the leading cause of transfer to the Sick Baby Unit (SBU) and the largest single cause of death among infants dying in the SBU (unpublished, Nyunt Nyunt Wai). Nevertheless a literature search based only on MEDLINE failed to find articles dealing with nosocomial neonatal diarrhea in developing countries. We therefore initiated this retrospective study to describe certain features of the problem.

Methods

The study was carried out at Central Women's Hospital, the largest of the hospitals serving maternity patients in Rangoon, Burma with about 16,000 deliveries annually. Formerly, newborn infants were cared for in nurseries separated from the mothers, and mothers and infants remained in hospital routinely for four to five days. During the period of the study, policies were introduced to promote "early discharge" (routine discharge one or two days after delivery) and "rooming in." Infants were said have "rooming in" if they were cared for by their mothers and stayed in cribs at the mother's bedside in large wards rather than residing in ward nurseries separated from their mothers except for feeding time. For infants born by cesarean section, these changes in policy did not apply and infants were for the most part cared for in nurseries by nurses.

This analysis is based on a review of the log from the SBU. Data include all admissions to the unit between 1981 and 1986 with diarrhea as the admitting diagnosis. The total numbers and types of delivery, obtained from annual hospital records, were used as denominators in calculating rates. Using the monthly totals which were available for the years 1983 through 1986, an estimate was made of the monthly distribution of deliveries for the period 1981–86.

Address reprint requests to David A. Sack, MD, Division of Geographic Medicine, Department of International Health, Room 5505, Johns Hopkins University, School of Hygiene and Public Health, 615 N. Wolfe Street, Baltimore, MD 21205. Co-authors Aye and Thwe are with the Bacteriology Research Division, Department of Medical Research, Rangoon, Burma; Dr. Wachsmuth is with the Enteric Bacteriology Laboratory, Centers for Disease Control, Atlanta, GA; Dr. Kyi is with Central Women's Hospital, Rangoon. This paper, submitted to the Journal December 14, 1989, was revised and accepted for publication October 30, 1990.

Editor's Note: See also related editorial p. 415 this issue.

TABLE 1—Numbers and Rates of Neonatal Diarrhea (transfers to Sick Baby Unit) Occurring in Infants Born by Vaginal Delivery or Cesarean Section by Year

Year	Vaginal (rate)	Cesarean (rate)	Overall rate
1981	236 (16.9)	133 (65.5)	22.4
1982	202 (14.4)	106 (50.1)	18.6
1983	130 (8.8)	63 (31.4)	11.2
1984	127 (8.9)	125 (54.6)	14.9
1985	92 (6.4)	123 (53.6)	12.8
1986	94 (6.6)	109 (49.5)	12.7
Totals	881 (10.3)	659 (51.0)	15.4

*Rate in cases per thousand live births

Episodes of mild diarrhea in which the infant was not transferred to the SBU would not have been detected in this surveillance, nor would episodes occurring after discharge if the child did not return to the same hospital. This review therefore underestimates the numbers of cases actually occurring, but there was no apparent bias over time since the record-keeping and referral system remained constant.

Data analysis was performed using Crunch Interactive Statistical Package Version 3.0 with a PC DOS microcomputer. Statistical tests used were chi square, Fisher's exact or Mann Whitney tests where appropriate.

Results

During the six years under study, 1,540 infants were admitted to the SBU because of diarrhea. The annual number of cases ranged from 193 to 368 (rates of 11.2 to 22.4 per 1,000 live births). Overall rates for the first two years were higher than for subsequent years (Table 1). The monthly rates were higher during the post monsoon dry season (rates averaged 17.5 per 1,000 live births, October–January) and lower during the hot monsoon season (rates averaged 13.5 per 1,000 live births for May–August).

The overall yearly rates decreased during the period of study associated with a decrease in vaginal delivery rates. By contrast the rates associated with cesarean section were higher and remained unchanged (Table 1). Overall, infants born by section had a five times higher risk of developing diarrhea than those born by vaginal delivery (95% CI = 4.5, 5.5).

One hundred eighty-four (12 percent) of the infants with neonatal diarrhea died. Fatal cases tended to be older when admitted to the SBU (median of seven days for fatal cases vs six days for nonfatal, $p < .001$), to be smaller at birth (relative risk for birthweight < 2.3 kg = 4.6 (C.I. = 3.5, 6.0)). Infants with larger birthweights (> 3.6 kg) were also at higher risk (relative risk = 2.7 (C.I. = 1.2, 5.8, Fisher's exact = 0.036) when compared to those with birthweights between 2.3 and 3.6 kg. Although infants born by section were at a higher risk of diarrhea, they were not at a higher risk of dying with diarrhea. Case fatality rates were not different according to sex or ethnic group, but infants born of mothers with more than seven previous pregnancies were at higher risk ($p = 0.02$).

The available records did not permit a comparison of the case fatality rates of different clinical syndromes or of the possible benefits of specific forms of treatment on the case fatality rates.

Discussion

This study has demonstrated that diarrhea sufficiently severe to lead to transfer to the SBU was common in this Rangoon hospital between 1981 and 1986. The case fatality rate of 12 percent reinforces the severity of the problem. Cesarean section was the most prominent risk factor for diarrhea identified in the study (RR = 5.0). These infants tend to stay in the hospital longer, to be handled more by hospital staff and less by mothers, and to be fed more formula and less breast milk than other infants.

For children born by vaginal delivery, the rates of neonatal diarrhea decreased three-fold during the period of study. This decrease may have been related to "rooming in" and "early discharge" policies; however, we were not able to determine if these policies were effective, since the changes were phased in over a prolonged period and did not allow for a "before–after" analysis.

There are obvious limitations in the study which need to be addressed. As a retrospective study the cases detected likely underestimate the problem. Case definitions depended on the diagnosis of "diarrhea" at the time of transfer, rather than an actual description and quantitation of symptoms. Cases were not evaluated bacteriologically to determine the causative agents. Finally, there may be other relevant risk factors (types of formula, hand washing, etc.) which were not evaluated.

Nevertheless, the study clearly documents the importance and endemicity of the problem and suggests the need for further research as well as control strategies based on observations of current practice. ☐

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Effect of Hospital Utilization Review on Medical Expenditures in Selected Diagnostic Areas: An Exploratory Study

Thomas M. Wickizer, PhD

ABSTRACT

Quarterly claims data on 43 insured groups were analyzed through multivariate techniques to explore whether the effects of hospital inpatient utilization review vary across selected broad diagnostic areas. Findings suggest that utilization review was associated with decreases in expenditures of approximately 15 percent for diagnoses within the surgical area, a lesser decrease within the mental health area, and still lesser decrease within the medical area. However, these measurements are imprecise both because of the small numbers and the aggregated diagnoses in each category. (*Am J Public Health* 1991;81:482–484)

Introduction

Hospital inpatient utilization review (UR) has assumed a prominent role as a private cost containment activity[1,2]; 65 percent of group insurance plans now operate under some form of UR.[3] UR evaluative research has produced mixed results,[4] with some studies finding large reductions in hospital use associated with UR[5–10] and other studies finding little if any effect.[11–15] Unfortunately, many of the UR evaluations suffer from methodological problems that raise questions about the reliability of the findings.[4,16]

One recently initiated study, conducted by the present author and colleagues, used multivariate techniques to analyze claims data on 223 insured groups. The study found UR was associated with significant reductions in hospital admissions, inpatient expenditures, and total medical expenditures.[17,18] However, the question of whether UR may have different effects across diagnostic categories was not examined, nor has this question been examined by other studies. The purpose of this paper is to report the findings of such an analysis.

Methods

The UR program analyzed, described in more detail elsewhere,[19] was established in 1983 by CNA Insurance Companies of Chicago and made available to policyholders as a benefit plan option. It consists of preadmission authorization and concurrent review and is compulsory for all employees and dependents. Patients failing to comply with UR procedures are subject to financial penalties. Physicians and hospitals are not subject to penalties.

The unit of observation for the analysis is the insured group. Expenditure data covering the period 1984 through 1986 were collected for selected diagnostic categories on 43 groups that adopted UR after operating without any cost containment program. Well-distributed geographically and representative of the under-65 population in terms of hospital use, these groups were selected for study because they provide longitudinal data (pre- and post-UR data) that permit the use of an efficient method of analysis. (Because of resource limitations, it was not possible to collect claims data by diagnosis for all 223 groups included in the larger study.) The average group includes approximately 1,300 insureds, comprising 550 employees and 750 dependents.

Quarterly claims data on the 43 groups were pooled, making 443 usable and complete observations available for analysis. To examine the effects of UR across diagnostic areas, measures were created representing mental health services, medical services, and surgical services. These measures were constructed by summing expenditures over selected diagnostic categories for which the insurance carrier routinely collected claims data. The term diagnostic area is somewhat imprecise. In general, it represents a set of conditions for several related ICD-9-CM diagnostic categories, but it bears no direct relationship to diagnosis related groups (DRGs) or other diagnostic classifications.

For the *mental health area*, expenditures were summed over alcohol, drug abuse, psychoses, and neuroses. For the *medical area*, expenditures were summed over diabetes, hypertension, infectious and allergy diseases, obstructive pulmonary disease, viral infections, and cutaneous diseases of the skin. For *surgical services*, expenditures were summed over benign neoplasms, diseases of the esophagus and stomach, diseases of the bones and cartilage, and musculoskeletal injuries.

Address reprint requests to Thomas M. Wickizer, PhD, Department of Health Services, SC-37, University of Washington, Seattle, WA 98195. This paper, submitted to the Journal March 13, 1990, was revised and accepted for publication August 27, 1990.

Descriptive information on the expenditure measures is presented in Table 1. The measures, which capture all inpatient as well as outpatient covered charges, represent services accounting for approximately 40 percent of total nonobstetric expenditures. Surgical services have the largest proportion of cases (28 percent) with quarterly expenditures over $30 per insured person. Included in the analysis is a set of 16 covariate factors, described in detail elsewhere,[19] representing employee population characteristics, health care market factors, and benefit plan features (see Appendix).

UR is measured by a binary variable that takes on a value of 1 if a group operated under UR during a given quarter, and 0 otherwise. It was not possible to distinguish preadmission authorization from concurrent review for purposes of analysis.

Ordinary least squares (OLS) regression was used to estimate a covariance model, constructed by differencing each observation from its time mean.[19–21] This model, which uses within-group (over time) variance to identify parameters, produces the same estimates as the standard fixed-effects (or least squares dummy variable) model but does not require inclusion of group-specific dummy terms, thereby saving degrees of freedom and improving computational efficiency. Since aggregate data collected over time were being analyzed on groups differing in size, procedures were performed to test for heteroscedasticity[22] and serial correlation.[23] The model's error terms were found to be heteroscedastic, and a standard weighting procedure was performed on the data to correct the standard errors. Specification tests revealed no evidence of serial correlation.

Three equations were estimated, one for each of the three services, with the covariate factors and the UR term included on the right hand side of the equations, in addition to a set of quarter dummy variables included to control for unmeasured temporal factors.

Results

The results of the analysis are presented in Table 2. As indicated, UR was associated with somewhat lower expenditures ($p < .07$) in the surgical area; after adopting UR the average group's expenditures for this diagnostic area declined by approximately 15 percent. This estimated percentage reduction implies an absolute reduction in surgical expenditures of ap-

TABLE 1—Descriptive Information on Dependent Measures (N = 433)

Measures	<$10.00 per Insured Person per Quarter[a]	$10.00 to $19.99 per Insured Person per Quarter	$20.00 to $29.99 per Insured Person per Quarter	>$30.00 per Insured Person per Quarter
Expenditures on Medical Services[b] (%)	109 (25)	184 (43)	88 (20)	52 (12)
Expenditures on Surgical Services (%)	57 (13)	148 (34)	106 (25)	122 (28)
Expenditures on Mental Health Services (%)	226 (52)	76 (18)	60 (14)	71 (16)

a) Expenditure data were adjusted by the Medical Care Price Index (MCPI) to reflect January 1985 prices.
b) The means and standard deviations (SD) for the three measures are: 1) Medical Expenditures, mean $17.78, SD $12.06; 2) Surgical Expenditures, mean $25.55, SD $21.32; and 3) Mental Health Expenditures, mean $16.16, SD $20.83.

TABLE 2—The Effects of UR on Expenditures for Three Diagnostic Areas: Mental Health, Medical, and Surgical Services (N = 433)

Measures[a]	Medical Expenditures	Surgical Expenditures	Mental Health Expenditures
Intercept	0.142* (0.087)[b]	0.131 (0.095)	0.261 (0.244)
% Female Employees under Age 30[c]	0.100 (0.173)	0.134 (0.193)	−0.204 (0.489)
% Male Employees Ages 30 to 50	0.023 (0.046)	0.058 (0.051)	0.157 (0.131)
% Female Employees Ages 30 to 50	0.005 (0.137)	−0.105 (0.152)	0.428 (0.386)
% Male Employees over Age 50	0.012 (0.029)	0.024 (0.032)	−0.015 (0.081)
% Female Employees over Age 50	−0.155 (0.300)	0.033 (0.332)	−0.597 (0.845)
% Childbirth/Pregnancy Expenditures	−0.901* (0.563)	−0.277 (0.623)	−0.086 (1.585)
% Heart Disease Expenditures	−0.298 (0.635)	−0.164 (0.703)	0.648 (1.789)
HMO Penetration Rate	0.015 (0.026)	0.037 (0.028)	−0.015 (0.072)
Office-Based Physicians per 1,000 Population	1.265 (0.917)	−0.822 (1.014)	−1.466 (2.584)
% General Practice Physicians	−2.246 (1.956)	−1.397 (2.164)	−0.841 (5.511)
Hospital Occupancy	−0.028** (0.013)	−0.029* (0.015)	−0.038 (0.037)
% Charges Covered for Semi-Private Room	0.001 (0.002)	−0.001 (0.003)	0.010 (0.007)
% Charges Covered for Hospital Outpatient Services	0.004 (0.003)	−0.000 (0.004)	−0.021** (0.010)
% Charges Covered for Physician Office Services	0.002 (0.005)	−0.001 (0.006)	−0.014 (0.016)
Deductible	−0.014 (0.017)	0.034* (0.019)	0.084* (0.049)
Utilization Review[d]	−0.062 (0.097)	−0.157* (0.108)	−0.146 (0.274)

NOTE: R^2 is not shown because this does not provide a meaningful indication of the model's explanatory power for mean-differentiated equations.
a) The coefficients of the quarter dummy variables are not included in Table 2 for economy of presentation.
b) Standard errors are shown in parentheses.
c) The reference variable is % Male Employees under age 30.
d) Since the dependent variable is measured in logarithms, the coefficients of the utilization review (dummy) variables can be interpreted as approximate percentage differences. The actual percentage difference in expenditures pre- versus post-UR is slightly lower than indicated by the coefficients. For example, the coefficient −0.157 for the surgical expenditure equation implies a 14.7 percent reduction in expenditures.
*$p < .10$; **$p < .05$

proximately $17.25 per insured person per year. Although the estimated UR coefficient for the mental health area is almost as large, it is not statistically significant due to the large standard error estimate. It appears UR has a less important effect on medical expenditures. As Table 2 shows, with few exceptions the covariate factors are not statistically significant.

Discussion

This exploratory analysis has sought to contribute to an improved understanding of an important question regarding the effects of hospital UR. The findings suggest UR's effects vary across diagnostic areas, and that UR may be more effective in reducing expenditures on diagnoses likely to involve surgical services than other services, especially medical services. One plausible explanation for this finding is that UR protocols may be better developed for surgical cases than for other cases. To work effectively, UR must have clinical protocols that can be followed and used to establish the need for hospital care. Such protocols may be better developed for surgery than other areas, which may explain, in part, why UR appears to be more effective for cases likely to involve surgical services. However, given the limitations of the data, these findings should be considered preliminary. Future research should conduct analyses using more complete and detailed data to examine this question in greater depth. ☐

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APPENDIX—Descriptive Information on Covariates Included in the Analysis (N = 433)

Measures	Mean (Standard Deviation) %
% Female Employees under Age 30	10.1 (8.0)
% Male Employees under Age 30	11.9 (6.8)
% Male Employees Ages 30 to 50	35.2 (9.2)
% Female Employees Ages 30 to 50	18.6 (9.8)
% Male Employees over Age 50	17.3 (10.0)
% Female Employees over Age 50	6.5 (3.7)
% Childbirth/Pregnancy Expenditures[a]	8.1 (7.4)
% Heart Disease Expenditures[a]	4.1 (5.6)
HMO Penetration Rate	11.5 (13.7)
Office-based Physicians	2.3 (1.2)
% General Practice Physicians	13.4 (11.2)
Hospital Occupancy	67.2 (8.7)
Deductible	$130 ($52)
% Charges Covered for Semi-Private Room	86.5 (24.3)
% Charges Covered for Physician Office Services	83.7 (16.4)
% Charges Covered for Hospital Outpatient Services	88.2 (20.4)
Utilization Review	0.75 (0.44)

a) Measured as a percentage of total expenditures.

Tracking Progress toward National Health Objectives in the Elderly: What Do Restricted Activity Days Signify?

Delia Scholes, PhD, Andrea Z. LaCroix, PhD, Edward H. Wagner, MD, MPH, Louis C. Grothaus, MA, and Julia A. Hecht, PhD

ABSTRACT

Restricted activity days is the measure by which the 1990 health objectives for prevention of functional disability in older adults will be evaluated. Yet its significance in older populations is poorly understood. We evaluated its use as an outcome measure for a randomized trial designed to impact upon physical function in elderly HMO enrollees. As predicted, restricted activity days was more correlated with physical disability measures than with other health status measures. Distributional properties and rates of missing data were shortcomings. (*Am J Public Health* 1991;81:485–488)

Introduction

In 1979, the US Surgeon General's report[1] designated the prevention of functional disability as the principal goal of health promotion in older adults. Restricted activity days (RAD), a standard item on the National Health Interview Survey, was chosen as the primary measure of disability to be monitored with the goal of achieving a 20 percent reduction by 1990. Evaluation of progress toward this goal is hampered because the meaning of RAD as a measure of health status in the elderly remains uncertain.[2,3] In younger age groups, RAD reflect transient loss of ability to perform social roles at work, home, or school, usually because of a physical limitation. In the elderly, many of whom are retired, usual activities are more variable. For this reason, several investigators have called for a better understanding of the appropriateness and usefulness in the elderly of such traditional health status measures as RAD.[4-6]

Because of the Surgeon General's 1990 objectives, we chose RAD as a primary outcome measure for an ongoing randomized controlled trial of health promotion involving nearly 2,300 older adults. We hypothesized that RAD would be more strongly correlated with measures of physical disability than with measures of mental, social, and global self-perceived health. Since there has been some speculation that days spent in bed might be a better measure of health status than RAD,[7] we compared the associations of both variables with these other health status measures.

Methods

The study setting is Group Health Cooperative (GHC) of Puget Sound, an HMO (health maintenance organization) located in western Washington State. At the study's inception in 1987, approximately 11 percent of enrollees were age 65 or older. This report is based on data from 2,289 enrollees, 36 percent of a random sample of 6,328 seniors, who returned completed baseline questionnaires and agreed to participate in the trial.

The questionnaire covered a broad range of health beliefs, behaviors, and conditions and included a number of standard single-item measures and multi-item scales. Information was obtained on the number of RAD and days spent in bed for the preceding 12 months using the questions described in Table 1, Section A.

For comparison with RAD and bed days we selected a number of baseline measures of physical health (Table 1, Section B), several measures of mental and social health (Table 1, Section C), and two self-evaluations of health status (Table 1, Section D). Spearman rank correlation coefficients were used to assess the association between the disability day variables and the other health status measures. Measures within each health status domain (physical, psychosocial, and self-evaluated health) were correlated with each other and within-group correlations were compared with correlations of each measure with disability days.

To explore reasons for non-participation in the trial, we interviewed a random sample of 175 non-participants by telephone using an abbreviated version of the same questionnaire. Nonparticipants were similar to participants with respect to age, gender, the prevalence of chronic disease, and number of RAD. They had less education, lower incomes, higher rates of smoking, worse self-perceived health, and less participation in social activities.*

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*Wagner EH, Grothaus LC, Hecht JA, LaCroix AZ: Factors affecting participation in a senior health promotion trial. Manuscript under review.

From the Center for Health Studies, Group Health Cooperative of Puget Sound. Address reprint requests to Delia Scholes, PhD, Center for Health Studies, Group Health Cooperative of Puget Sound, 1730 Minor Avenue, Suite 1600, Seattle, WA 98101. This paper, submitted to the Journal April 11, 1990, was revised and accepted for publication September 11, 1990.

TABLE 1—Selected Health Status Indices from Senior Baseline Questionnaire

Health Status Measure	Description and/or Item Content	Source
A. Disability Days		
1. Restricted activity days	Two items: "In the past 12 months, did you cut down the things you usually do, such as going to work or working around the house, because of illness or injury?" IF YES: "How many days did you cut down on the things you usually do because of illness or injury?"	Adapted from NHIS[8]
2. Days spent in bed	Two items: "In the past 12 months, did you ever stay in bed because of an illness or injury?" IF YES: "How many days did you stay in bed at least half the day because of illness or injury?"	Adapted from NHIS[8]
B. Physical Health Measures		
1. Functional limitations	Guttman scale of seven daily living activities: "For how long (if at all) has your health limited you in each of the following activities? (1) The kinds or amounts of vigorous activities you can do . . . [least impaired] . . . (7) Eating, dressing, bathing or using the toilet [most impaired]". Coefficient of reproducibility = .91, coefficient of scalability = .61.	Adapted from Ware[9,10]
2. Illness	Single item: "In the past 12 months, have you had a major personal illness?"	
3. Hospitalization	Single item: "In the past 12 months, were you ever in the hospital overnight for physical health problems?"	
4. Pain	Single item: "We would like to find out about any physical pain you may experience in your daily life. Below is a scale from 0 (no pain) to 10 (very severe pain). Please place an 'X' on the line in the spot that most closely represents how much pain you experience on a daily basis."	Ware[11]
5. Symptoms	13 questions about a range of symptoms experienced in the past week: e.g. "'I felt dizzy or lightheaded when I sat up or stood up"; "I had stomach cramps."	Developed for the study
6. Role limitations	Seven item scale on short-term physical limitations: "During the past 4 weeks, have you had any of the following problems with your work or other daily activities as a result of your physical health? a. Were limited in the kind of work or other activities you usually do?; g. Required special assistance (the assistance of others or special devices) to perform the activities you usually do."	Ware[10]
7. Heart Trouble	Single item: "Here is a list of medical conditions which usually last for some time. Have you had any of these conditions in the past 12 months?" Affirmative to "Heart trouble".	Belloc[12]
C. Social/Psychological Measures		
1. Life Satisfaction	11-item scale: "Below each question is a scale from 0–10, where 0 is "very satisfied" and 10 is "very dissatisfied". Please place an 'X' on the line by the number that most closely represents your response to each question."	Patrick[13]
2. Positive Affect/Pep and Vitality	Combined score from 10-item Positive Affect subscale and a number of items from a second subscale (Pep and Vitality): "During the past month . . . (a) how much of the time have you felt that the future looks hopeful and promising? . . . (m) how much energy, pep or vitality did you have or feel?" (6 response categories)	Adapted from Veit[14]
3. CES-D	10-item adaptation of the Center for Epidemiological Studies Depression Scale. "Please indicate by circling the appropriate response how often you have felt this way during the past week: (a) I was bothered by things that usually don't bother me . . . (j) I could not 'get going'."	Adapted from Radloff[15]
4. Stress	Single item: "Pictured below is a scale from 1 (none) to 10 (a great deal). Which number on the scale best represents how much stress you have been under for the past year?	Bailey[16]
5. Social Support	Single item: "Do you belong to a close circle of friends, a group of people who keep in touch with each other?"	Morgan[17]
6. Nervous Condition	Single item: "Here is a list of medical conditions which usually last for some time. Have you had any of these conditions in the past 12 months?" Affirmative to "Chronic nervous or emotional problems".	
D. Self-Evaluations of Health		
1. Self Perceived Health	Single item: "Would you say, in general, your health is (1) excellent . . . (5) poor?"	Ware[10]
2. Current Health Outlook	9-item subscale of General Health Perceptions Index: e.g. "My health is excellent"; "I have been feeling bad lately." (5 response categories)	Ware[10]

Results

Missing Values

In the 2,289 participants with completed baseline questionnaires, 11 percent of responses on RAD and 4 percent on bed days were missing. Missing data were more frequent for both measures among older respondents and those with poorer self-evaluated health (Figure 1.a., 1.b.).

Baseline Distribution of RAD and Bed Days

Both disability day variables had highly skewed distributions, as has been found for many measures of physical health in community-dwelling older populations.[14] Seventy-four percent of respondents reported no RAD and 73 percent reported no days spent in bed in the past 12 months.

Other Health Status Measures

Spearman correlations of RAD and bed disability days with selected measures of physical, psychosocial, and self-evaluated health status are presented in Table 2. Restricted activity days and bed days were highly correlated (.61); when the subset of RAD due to bed days was excluded, the correlation was .53. As hypothesized, the number of RAD was most strongly correlated with indicators of physical health (correlations ranged from .20 to .48), including the other two disability measures, functional and role limitations (r's = .34 and .39, respectively). Within-group correlations among physical health status measures ranged from .11 to .60.

Moderate correlations between RAD and the two self-evaluations of health were also found: .24 for self-reported health and .30 for current health outlook. However, these two measures were more highly correlated with each other (.68). The weakest associations with RAD were for the social support and mental health measures. Restricted activity days was more strongly correlated than bed days with all the physical health measures, except for hospitalization in the past 12 months (Table 2).

Our assessment of correlations between RAD and a subset of these variables in the sample of nonparticipants showed findings similar to the larger sample; RAD continued to be more strongly correlated with the available physical health status measures and had stronger correlations than bed days with most health status measures.

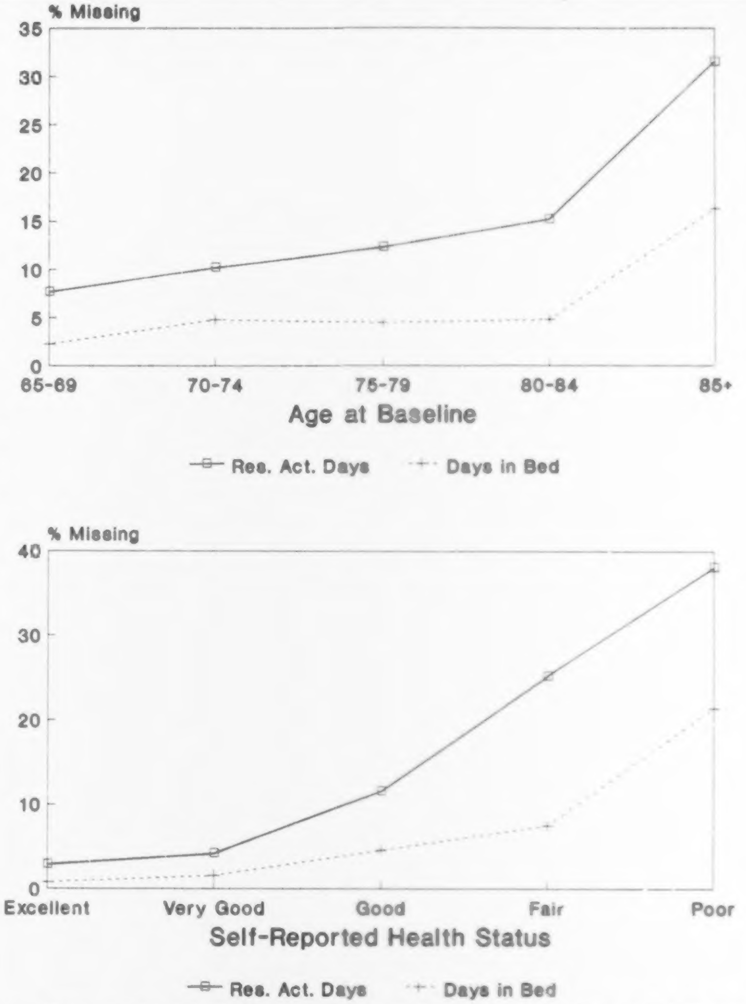

FIGURE 1—Rate of Missing Data for Restricted Activity Days and Days in Bed by Age Group and Self-reported Health Status

Discussion

The vague nature of RAD has been cited as a drawback to its use.[7] Our findings indicate that the potential shortcomings do not preclude its use in studies where physical disability is the health dimension of interest. The validity of using RAD is supported by its being most highly correlated with measures of physical health status and with functional limitation measures in particular.

Our results do not support the suggestion that bed days may be a better indicator of physical disability; its correlations with virtually all other health status measures were weaker than those for RAD.

While these results support our initial hypothesis, several limitations to the use of RAD deserve mention. Both disability days variables had very skewed distributions in this sample of essentially well elderly, limiting their usefulness in describing the full range of physical health status. In addition, over 10 percent of the data on RAD were selectively missing, an indication that this variable, in self-administered form, posed problems for participants.

The proportion of the sample that agreed to participate in this study represents a potential limitation of this analysis. This would be particularly so if non-participants were more disabled than those who agreed to be part of the study. Our study of nonparticipants allowed us to replicate this analysis in part and provided information that they had similar levels of disability and similar correlation patterns.

As evaluation of the 1990 national

TABLE 2—Correlation Matrix[1,2] for Disability Days, Physical and Social/Psychological Health Indices, and Self-Evaluations of Health

	A.		B.							C.						D.	
	1	2	1	2	3	4	5	6	7	1	2	3	4	5	6	1	2
A. Disability Days																	
1. Restricted Activity Days	/																
2. Days spent in bed	.61	/															
B. Physical Health Status Measures																	
1. Functional Limitations Scale	.34	.22	/														
2. Illness in past 12 months	.48	.46	.30	/													
3. Hospitalization in past 12 months	.41	.51	.22	.60	/												
4. Pain scale	.24	.15	.47	.17	.11	/											
5. Symptoms	.22	.13	.35	.17	.11	.34	/										
6. Role limitations	.39	.23	.56	.28	.19	.42	.42	/									
7. Heart disease	.20	.15	.25	.32	.28	.14	.21	.19	/								
C. Social/Psychological Measures																	
1. Perceived quality of life	.17	.10								/							
2. Positive affect	.20	.13								.60	/						
3. CES-D Depression Scale	.17	.12								.46	.59	/					
4. Stress	.18	.13								.46	.51	.39	/				
5. Circle of friends	.02	.05								.18	.19	.09	.06	/			
6. Nervous/emotional problems	.13	.05								.19	.25	.30	.30	.05	/		
D. Measures of General Well-Being																	
1. Self-reported health	.24	.16														/	
2. Current health outlook	.30	.21														.68	/

[1] For ease of presentation, all measures of health status are scored so that high scores indicate poorer condition (e.g. poorer life satisfaction, poorer affect, more depressive symptoms, etc.).

[2] Upper and lower limits for 95% confidence intervals for correlations can be obtained by adding and subtracting the following quantities to the correlations in the table: +/−.044 for correlations between 0 and .30; +/−.039 for correlations of .31 to .45; and +/−.033 for correlations of .46 to .61 For correlations involving the CES-D, corresponding figures are: .053, .047, and .039, due to higher rates of missing data.

health objectives continues, these findings may help to better characterize the role of RAD in older adults and provide added perspective on their place in the more broadly defined health objectives for the elderly for the year 2000. □

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Risk Communication Activities ~~of~~ State Health Agencies

Caron Chess, MS, Kandice L. Salomone, MS, and Peter M. Sandman, PhD

ABSTRACT

Surveys concerning the risk communication practices and needs of state health agencies were completed by agency commissioners and designated staff of 48 states and territories. These data indicate that agencies are expending more effort on responding to requests for information than on initiating dialogues with interested constituencies or alerting the public to risk. The data also suggest a gap between the stated philosophy and practice of the agencies. (*Am J Public Health* 1991; 81:489–491)

Introduction

Increasingly, government agencies have been concerned about how to communicate environmental health risks to the public. There is a growing literature prescribing how agencies should practice risk communication,[1–4] describing case studies of agency efforts,[5–9] and investigating public perception of risk.[10–13]

Nevertheless, little is known about the risk communication activities and needs of public health agencies. We report briefly the results of a national survey of state health agencies. Two detailed reports, one dealing with the data's implications for risk communication training and the other dealing more fully with analyses of state health agencies' risk communication activities, are available from the authors.

Methods

Early in 1989, the Association of State and Territorial Health Officials (ASTHO) distributed a risk communication survey (developed by the Rutgers Environmental Communication Research Program) to the health commissioners of the 50 states, five territories, and the District of Columbia. The survey was developed primarily to assess health agencies' risk communication needs. Therefore, the questions focused solely on risk communication rather than on the broader context of respondents' responsibilities, priorities, and resources. Health commissioners were asked to complete one survey themselves and to designate one person with technical background and one with communication background to complete individual copies of a similar survey. The surveys elicited returns from 128 respondents from 48 states and territories, with the 85 staff respondents accounting for 66.4 percent of all data analyzed and 43 commissioners accounting for 33.6 percent of analyzed data. Returned surveys reflect a 76.8 percent response rate for commissioners and a 75.9 percent response rate for agency staff.

Results

According to staff respondents, the five environmental health issues on which agencies devoted the most time to communicating with the public in the past year were: drinking water contamination (90.6 percent); radon in indoor air (64.7 percent); potential cancer clusters (62.4 percent); clean up of hazardous waste sites (57.6 percent); and pesticides (43.5 percent).

The dominant risk communication activity of agency staff is responding to inquiries (mean percentage of time = 62.1 percent). The remaining time is split between designing and planning outreach activities to alert the public to environmental health risks (21.8% percent) and soliciting input from the public on agency policies and actions (12.7 percent).

Staff reported that their agencies communicate on a daily basis with individuals who contact them with questions about environmental issues, but they initiate contact only about once a week. They talk only once or twice a month with environmental advocacy groups, other groups that address specific environmental issues, and neighborhood community groups.

When responding to inquiries, staff reported that the people with whom they speak most frequently have learned about a potential risk from the media and that they deal most frequently with people who have emotional responses to risk (Table 1).

When soliciting input from people outside the agency on agency policies and environmental health issues, staff reported that their agencies used telephone contacts and informal meetings more frequently than seven other sources of input

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Address reprint requests to Caron Chess, MS, Associate Director, Environmental Communication Research Program, Cook College, Box 231, Rutgers University, New Brunswick, NJ 08903–0231. Ms. Salomone is Research Methodologist at ECRP and Dr. Sandman is ECRP Director. This paper, submitted to the Journal December 20, 1989, was revised and accepted for publication August 13, 1990.

TABLE 1—Average Responses From Staff of Inquiry Sources, Inquiry Type, Sources of Outside Input, and Outreach Methods

	Mean[a]	Standard Deviation[a]	Median[a]
Inquiry Source			
Individuals who learned about risk from the media	3.55	1.06	4.00
Individuals from environmental or health advocacy groups	2.41	1.02	2.00
Individuals who learned about risk from agency outreach methods	2.25	1.05	2.00
Individuals who raise concerns not yet known to agency	2.11	1.13	2.00
Inquiry Type			
Inquiring people have emotional responses to risk	3.62	1.23	4.00
Inquiring people are concerned about other issues besides risk	3.46	1.16	3.00
Inquiring people want explanations of risk data	2.87	1.33	3.00
Inquiring people want information about agency's risk management process	1.77	1.03	1.00
Source of Outside Input			
Telephone contacts	2.66	1.32	3.00
Informal meetings	2.45	0.95	2.00
Public hearings/other formal public meetings	2.05	1.00	2.00
Formal task forces	1.94	1.05	2.00
Hotlines	1.93	1.40	1.00
Public announcements requesting responses to written proposals	1.77	1.00	1.00
Questionnaires	1.33	0.57	1.00
Polls	1.07	0.34	1.00
Outreach Methods			
Telephone calls to individuals	2.65	1.42	3.00
Via local officials	2.44	1.10	2.00
Contact by technicians during sampling	2.33	1.40	2.00
Agency attendance at meetings	2.28	0.88	2.00
Health advisories	2.17	1.07	2.00
Initiating media stories	2.10	1.12	2.00
Hearings/other formal public meetings	1.95	0.85	2.00
Flyers/handouts	1.90	1.07	2.00
Public service announcements/ advertisements in the media	1.89	1.07	2.00
Via community or environmental groups	1.87	1.02	2.00
Agency-sponsored informal meetings	1.81	0.76	2.00
Drop-in sessions in the affected area	1.75	1.07	1.00
Posting of announcements	1.72	0.95	1.00
Mass mailings	1.52	0.85	1.00

[a] Descriptive statistics were calculated from a five-point Likert-type scale where (1) indicated an answer of "seldom or never," (2) indicated an answer of "monthly," (3) an answer of "weekly," (4) an answer of "daily," and (5) indicated an answer of "constantly."

(Table 1). However, no single approach was used very often—perhaps a few times a month, at most.

Much the same can be said for the frequency with which agencies conduct activities to alert the public to potential environmental hazards. Across 14 different methods, the most any one method was practiced was a few times a month. Most methods were used monthly and a few were used less than once a month (Table 1).

We probed agency philosophy with 10 statements. More than two-thirds of all respondents either "agreed" or "strongly agreed" with the following five statements:

- "communicating with the public about environmental risk is an important agency priority" (90.6 percent);
- "risk communication training is important to improve agency efforts" (87.3 percent);
- "a major barrier to effective environmental risk communication is public misunderstanding" (82.4 percent);
- "a major barrier to environmental risk communication is insufficient training for new and existing staff" (75.8 percent); and
- "a major barrier to effective environmental risk communication is insufficient staff for risk communication" (67.2 percent).

More than two-thirds of all respondents either "disagreed" or "strongly disagreed with these two statements:

- "communicating with the public about environmental issues does not make much of an impact" (88.2 percent); and
- "better to communicate less than to risk stirring up problems" (76.4%).

More than half "disagreed" or strongly disagreed with

- "the public should be encouraged to leave risk management decisions to the professionals" (57.1 percent).

Discussion

These data suggest that agencies, in spite of their philosophical commitment, are expending more effort on responding to requests for information than on initiating dialogues with interested constituencies or alerting the general public to risk.

Most agency efforts are devoted to responding to inquiries from the public, and most inquiries from the public are generated by the media. Although respondents indicated that their agencies spent more than 20 percent of their communication time on outreach activities, they reported using most outreach methods—including contacting the media—fairly infrequently (monthly or less), again suggesting that they may be reacting to communication events rather than shaping them.

The majority of respondents disagreed with the statement that the public should be encouraged to leave risk management decisions to the professionals, but no methods to solicit input were used with any frequency. This may indicate the flexibility that is important when dealing with a variety of audiences, or it may suggest that little time is spent in soliciting public input and even less in developing plans and policies on how to solicit input. ☐

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Capacity of US Labs to Provide TLI in Support of Early HIV-1 Intervention

Ronald O. Valdiserri, MD, MPH, G. David Cross, MS, A. Russell Gerber, MD, Robert E. Schwartz, BS, and Thomas L. Hearn, MS

ABSTRACT

We surveyed laboratories to assess their capacity to perform T-lymphocyte immunophenotyping. Of the 1026 respondents, 279 located in 41 states and the District of Columbia performed this type of testing. Most laboratories were located in hospitals, reported a low weekly test volume, and indicated that it took 6–24 weeks for flow cytometer operators to become proficient. Many laboratories appear to have the capacity to perform additional CD4+ cell testing, but training additional operators may be necessary. The paucity of laboratories performing T-lymphocyte immunophenotyping in the public sector may affect referral patterns from that setting. (*Am J Public Health* 1991;81:491–494)

Introduction

In 1989, the US Public Health Service recommended that all persons infected with human immunodeficiency virus type-1 (HIV-1) have CD4+ cell counts performed at least every six months.[1] This recommendation is based upon the knowledge that CD4+ cell counts are a major predictor of disease progression in HIV-1 infection,[2] and that information provided by this test is used to make therapeutic decisions about initiating antiretroviral therapy and prophylactic therapy for *Pneumocystis carinii* pneumonia.[3] Because CD4+ cell counts are currently performed by flow cytometric analysis, a technique that is not commonly used in many laboratories, a survey was conducted by the Division of Laboratory Systems, Public Health Practice Program Office (PHPPO) of the Centers for Disease Control (CDC) to learn more about the national availability of this testing and about testing practices in the laboratories currently providing CD4+ cell testing services.

Methods

In August 1989, 1,658 laboratories were surveyed to learn more about their T-lymphocyte immunophenotyping (TLI) testing practices; 1,348 were currently enrolled in a voluntary HIV-1 antibody testing performance evaluation program conducted by the CDC,[4] and 310 were enrolled in a flow cytometry proficiency testing program sponsored by the College of American Pathologists (CAP). The sur-

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From the Division of Laboratory Systems, Centers for Disease Control, Atlanta, GA. Address reprint requests to Ronald O. Valdiserri, MD, MPH, Director, Division of Laboratory Systems, Public Health Practice Program Office (G25), Centers for Disease Control, Atlanta, GA 30333. This paper, submitted to the Journal March 9, 1990, was revised and accepted for publication August 6, 1990.

TABLE 1—Percentage Distribution of Laboratories Performing and Not Performing TLI, according to Type of Laboratory

Type of Laboratory	% Performing TLI (N:279)	% Not Performing TLI (N:747)	Total % (N:1026)
Hospital, non-blood bank	60	28	37
Blood bank	12	33	27
Independent	15	18	17
Health Department	1	13	10
Research	7	2	3
All Other	5	6	6
Total	100	100	100

TABLE 2—Frequency of Weekly T-Lymphocyte Immunophenotyping Test Volume* by Laboratory Type

	Number (%) of Samples Tested Weekly			
	<50	50–199	200–500	>500
Hospital, non-blood bank (N=167)	117 (70)	41 (25)	—	9 (5)
Blood bank, hospital (N=26)	21 (81)	5 (19)	—	—
Blood bank, non-hospital (N=5)	3 (60)	2 (40)	—	—
Independent laboratory (N=42)	16 (38)	16 (38)	4 (10)	6 (14)
Research laboratory (N=20)	9 (45)	8 (40)	—	3 (15)
Health department (N=4)	3 (75)	1 (25)	—	—
Other/multiple response (N=9)	3 (33)	5 (56)	—	1 (11)
Total	172 (63)	78 (29)	4 (1)	19 (7)

*Does not include missing or incomplete data; therefore, N may be smaller than total sample.

veys, which were identical, were mailed independently. Laboratories that may have received one from both sources (i.e. enrolled in both programs) were requested to respond to the CDC survey only. Prior to data entry, obvious duplicate surveys (i.e. same laboratory name and address) were removed. After data entry, additional duplicate surveys were detected by examining for identical survey responses, across mailings, for laboratories which had the same postal zip codes; duplicate responses were removed from the data base. A total of 1,026 unduplicated surveys were available for analysis: 908 were from laboratories enrolled in the CDC-funded program and 118 were from laboratories which were enrolled in the CAP program, but not in the CDC program.

Results

TLI was performed by 279 (27 percent) of the laboratories, 116 from the CAP program and 163 from the CDC program. Information available on 734 of the 747 respondents not currently performing TLI testing indicated that 12 percent were considering providing this type of testing on-site within the next 12 months. Currently, 45 percent of the laboratories not performing TLI on-site offer the test through referral—most often (70 percent) to a commercial laboratory.

Table 1 shows the percentage distribution of laboratories performing and not performing TLI, by type, for the 1,026 laboratories returning surveys. Non-blood bank hospital laboratories (relative risk [RR] = 2.1; 95 percent confidence interval [95% CI]: 1.8, 2.5) and research laboratories (RR = 4.1; 95% CI: 2.1, 8.2) were more likely to perform TLI testing and health department laboratories (RR = 0.1; 95% CI: 0.04, 0.3) and blood banks (RR = 0.4; 95% CI: 0.3, 0.5) were less likely to do so.

TLI testing laboratories were located in 41 states and the District of Columbia, with the states of California, Texas, New York, Ohio, Pennsylvania, and Florida representing 39 percent of all TLI testing laboratories responding. Comparing the distribution of TLI testing laboratories by state to that state's AIDS rate for May 1989–April 1990,[5] a positive correlation was noted (Spearman rank correlation coefficient = .39, $p < 0.01$). Eighty-three percent of the laboratories reported that they tested specimens from an "off-site" location, i.e. not collected within the institution providing care.

In general, laboratories reported a low weekly testing volume (Table 2). Sixty-three percent of all respondents reported testing fewer than 50 samples per week and 38 percent of that group actually tested fewer than 10 samples weekly. High-volume testing (>500 samples per week) was noted in 15 percent (9/62) of independent and research laboratories—more frequently than in the 5 percent (10/211) of other types of laboratories (RR = 3.1; 95% CI: 1.3, 7.2).

Of those laboratories performing TLI testing, 82 percent had been testing for more than one year and 18 percent for one year or less. Not surprisingly, laboratories that have been performing TLI longer had higher weekly testing volumes. Forty percent (91/227) of laboratories that have performed TLI longer than one year tested 50 or more samples per week compared with 20 percent (10/49) of laboratories that have performed TLI one year or less (RR = 2.0; 95% CI: 1.1, 3.5).

When questioned about their routine sample staining protocol, some laboratories reported using more than one protocol, but most laboratories (75 percent) reported that their routine sample staining protocol was to stain whole blood samples, then lyse red blood cells. Eleven percent reported that they lyse red blood cells first and then stain samples, and 22 percent reported that they stain the sample after mononuclear cells have been separated by density-gradient centrifugation. When asked what cell markers were included in the normal battery of tests performed on each patient sample, the following observations were noted: CD4 (99%); CD8 (99%); CD3 (75%); CD2 (62%); CD19 (50%); CD20 (42%); CD45 (38%); DR (28%); CD14 (26%); NKH-1 (24%); CD57 (13%); 4B4 (11%); and Leu 19 (11%). The reported frequencies for other cell surface markers were less than 10 percent.

Self-reported quality assurance practices indicated that 97 percent of the laboratories had written protocols for specimen collection; 98 percent had protocols for routine calibration of their flow cytometers; and most (87 percent) performed instrument calibration daily. Ninety-one percent of laboratories performing TLI participated in an external proficiency testing program.

When questioned about the sources of training for persons performing TLI, the two most ~~frequent~~ types of training reported were instrument manufacturer's training school (84 percent) and in-house (83 percent). However, 48 percent of TLI laboratories reported taking advantage of courses offered by universities, professional societies, and industry (other than the instrument manufacturers). When questioned about the length of time needed for flow cytometer operators to become proficient in performing TLI testing, the responses range from one-to-five weeks to greater than 18 months—although 55 percent reported a range of six-to-24 weeks. Fifty-four percent (121/225) of the laboratories with longer than one year of TLI testing experience reported that four months or more were needed for the operator to become proficient, while only 31 percent (15/49) of the laboratories with a year or less testing experience reported that this same level of experience was required to successfully perform TLI (RR = 1.8; 95% CI: 1.1, 2.7).

Discussion

Early medical intervention in HIV-1 infection marks a new phase in our experience with the epidemic of acquired immunodeficiency syndrome (AIDS) in which efforts to prevent the spread of HIV-1 through counseling and testing will become integrated into medical treatment and case management. In addition to the obvious medical benefits of improving the health and longevity of infected persons, it has been suggested that early treatment programs targeted to persons who are already infected with HIV-1 might significantly augment current prevention efforts by providing further opportunities for behavioral intervention, thereby resulting in decreased viral transmission to others.[6]

To achieve the goal of providing appropriate and timely medical and psychosocial services to all individuals who are infected with HIV-1, it will be necessary to ensure that the capacity for service exists and that adequate delivery systems are in place. A 1989 policy analysis of the impact of Public Health Service recommendations for prophylactic treatment of *Pneumocystis* pneumonia indicated that "State, regional, county, and city governments will need to coordinate and plan" for the provision of health services required to support these recommendations.[7] This planning must encompass all relevant health care disciplines, including the laboratory which plays a critical role in the diagnosis and prognostic assessment of persons infected with HIV-1.

Because recommendations for both antiretroviral therapy and *Pneumocystis* prophylaxis are based upon the level of CD4+ cells in the peripheral blood, this laboratory test plays a critical role in early intervention efforts. The provision of CD4+ cell testing services to support early HIV-1 intervention efforts will require: adequate instrument capacities, adequate numbers of trained laboratory personnel, access (either direct or through referral) to CD4+ cell testing services by health care workers providing care to HIV-1 infected persons, including those at counseling and testing sites, and financial resources to support this service, especially for indigent persons who are HIV-1 infected.

This survey indicates that many laboratories across the nation are already performing CD4+ cell testing, and laboratories appear to have the capacity to perform additional testing. Obviously, our survey does not include all of the laboratories currently performing CD4+ cell testing. At the time this survey was conducted, we estimate that nearly 1,000 laboratories were performing CD4+ cell testing by flow cytometry (personal communication: Jeffrey Johnson, Program Resources Inc). Therefore, our survey sample probably represents at least 30 percent of all laboratories currently performing this test.

Assuming financial resources are available, the ability to perform additional TLI testing is a function of both instrument and personnel capacity. According to experts in flow cytometry, it is estimated that one fully-trained person should be able to test at least 10 specimens (six markers per specimen) within an 8-hour period. (Personal communication: Dr. Janet Nicholson, Centers for Disease Control.) Although our survey suggests that there may be equipment capacity to accommodate additional testing, it is likely that new personnel will have to be trained to provide for these services. As such, it is interesting to note that laboratories with more experience in performing TLI tended to give longer estimates of time to achieve flow cytometer operator proficiency than those laboratories which have had less experience in the field. There are several probable reasons for the observed range of responses. Experienced laboratories may have older equipment that required more training to learn how to operate compared to newer equipment which tends to be more "user friendly." Also, laboratories with less experience in TLI testing may underestimate the time that it takes operators to become truly proficient in this procedure. And finally, experienced operators may realize that, even with newer equipment, operator training is requred to ensure that the equipment is providing valid results.

These findings should not be interpreted to imply that adequate capacity exists to provide all of the TLI testing needed, only that instrument capacity exists to perform additional testing. The exact need for this testing is still not known, but will depend upon the number of HIV-1-infected persons and the stage of their illness. Untreated individuals who have been infected for longer periods will experience a more rapid decrease in the CD4+ count[8] and may therefore require TLI testing more frequently than twice yearly.

It is apparent from this survey that many of the laboratories which have capacity to perform additional TLI testing are not located in departments of public health. Therefore, bringing together counseling and testing services with medical treatment for early intervention will require a close coordination of services between the entry level of identifying persons who are infected with HIV-1, much of which takes place in the public sector, and the subsequent medical evaluation and provision of care which often takes place in the private sector. Efforts to prevent secondary morbidity and mortality by implementing early HIV-1 intervention programs will involve the assurance that reliable CD4+ cell testing services are available, either directly or through referral, to health care providers treating HIV-1 infected persons.

When HIV-1 counseling and testing takes place in a physician's office, individuals who are found to be infected can be further evaluated or referred for evaluation to determine whether the degree of their immunodeficiency, as determined by CD4+ cell testing, warrants antiretroviral therapy[9] and/or prophylactic therapy to prevent *Pneumocystis carinii* pneumonia.[1] However, considering that less than 4 percent of the publicly funded HIV-1 counseling and testing provided between January 1988 and September 1989 took place in private physician's offices and clinics,[10] it is unlikely that this scenario will apply to the majority of persons who are found to be infected with HIV-1.

Between January 1988 and September 1989, it is estimated that freestanding HIV-1 counseling and testing sites and clinics for the treatment of sexually transmitted diseases (STD) accounted for 65 percent of all publicly funded HIV-1 counseling and testing and 69 percent of all positive tests.[10] One way to ensure that HIV-1 infected persons identified by these programs have physical access to necessary CD4+ cell testing services is by developing a case-management system in which case managers are dedicated to coordinating and monitoring the referral of HIV-1 infected persons into systems which provide medical evaluation and, where appropriate, treatment.

Another option is to expand the provision of TLI testing in the public sector. Participants at the Fifth Consensus Conference on Testing for Human Retroviruses, hosted by the Association of State and Territorial Public Health Laboratory Directors, recommended that "State and local public health laboratories should be making plans immediately for provision of flow cytometric measurement of CD4+ cells" (Fifth Consensus Conference on Testing for Human Retroviruses, March 1990). Some states have already begun to explore this possibility,[7] although it may be practical only in those states with a large number of HIV-1 infected persons. Because storage time and temperature of the specimen may affect T-lymphocyte counts,[11] problems arising from specimen transport to distant sites may limit the degree to which this service can be centralized. Other states may opt to contract with independent commercial laboratories or with local laboratories in tertiary care facilities for the provision of high-quality TLI testing to HIV-1 infected persons.

Many of the concerns about provision of CD4+ cell testing would be minimized by simpler, less costly, tests. Although a number of other biologic markers have prognostic implications in persons infected with HIV-1, none to date have been shown to be superior to CD4+ cell enumeration.[2,12] Therefore, it is unlikely that CD4+ cell testing will soon be replaced as the major source of laboratory information about when to initiate therapy in persons with HIV-1 infection.

Regardless of the approaches taken, implementing early intervention activities by bridging the gap between identifying HIV-1 infected persons and providing medical care to these persons will involve the allocation of additional resources to provide the necessary laboratory services. It is unlikely that private insurance alone will be able to finance the laboratory services needed to achieve early intervention for all HIV-1 infected persons, thus public sector support may be needed. ☐

Acknowledgments

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Response of Hospitals, Skilled Nursing Facilities, and Home Health Agencies in Oregon to AIDS: Reports of Nursing Executives

Caroline McCoy White, DrPH, and Marie C. Berger, PhD

ABSTRACT

To describe experience and readiness to provide care for persons with acquired immunodeficiency syndrome (AIDS), responses to a mailed survey from 86 percent of chief nursing executives of the 207 hospitals, skilled nursing facilities (SNFs) and home health agencies (HHAs) in Oregon were summarized. As of January 1989, 44 percent reported having provided care; even in low incidence counties, 64 percent of hospitals had provided care. Adequate resources were reported by 79 percent in hospitals, 26 percent in SNFs, and 69 percent in HHAs. (*Am J Public Health* 1991;81:495–496)

Introduction

Developing resources for providing care to persons infected with the human immunodeficiency virus (HIV) or having acquired immunodeficiency syndrome (AIDS) is a societal concern. In Oregon, while the incidence of AIDS is not high on a per capita basis (697 cases reported by December 31, 1989 and a population of 2,687,000[1]), there is a significant and growing population for whom services must be provided. It challenges all health care institutions in Oregon.

Because nursing care is a critical component of the services that persons with AIDS require, we wanted to ascertain the extent to which agencies providing nursing services had cared for persons who had AIDS; and identify factors inhibiting care of AIDS patients. Chief nurse executives (CNEs) were selected as informants because they are responsible for developing policies and allocating resources for nursing care; they are also influential in the decisions made in their agencies.

Methods

In January 1989, we mailed questionnaires to CNEs of all hospitals (n = 79), skilled nursing facilities (whether or not hospital-associated, n = 66), and home health agencies (whether or not hospital-associated, n = 62) licensed and located in Oregon. The questionnaire probed the agency's current experience with providing care and resources to care for persons with AIDS; the overall impact of the epidemic on the agency; and the future role of the agency as perceived by the CNE. The questionnaire had both a prescribed format for response and opportunities for comments. After the second mailing and telephone follow-up, and eliminating the blank questionnaires, the over-all response rate was 86 percent. The CNEs whose responses were used for analysis represent 97 percent of hospitals, 76 percent of skilled nursing facilities, and 84 percent of home health agencies in Oregon. They include 81 percent of the 100 agencies (95 percent of hospitals, 77 percent of SNFs and 67 percent of HHAs) in high prevalence areas (defined as those counties where cumulative numbers of reported cases were 11 or more as of December 31, 1988[2]) and 92 percent of the 107 agencies (100 percent of hospitals, 74 percent of SNFs and 95 percent of HHAs) in low prevalence areas.

Results

Even though Oregon has a moderate prevalence rate for AIDS (251.6 cases per million population in 1989[1]), in early 1989 44 percent of the 207 agencies reported having provided care for persons with AIDS (Table 1). This includes 89 percent of the hospitals and 81 percent of HHAs in high incidence areas, but only 8 percent of the SNFs regardless of location. Even in low incidence counties, 64 percent of hospitals had provided care to persons with AIDS.

CNEs were asked to check specific agency resources for the care of AIDS patients and implementation of universal precautions (Table 2). As a group, CNEs of hospitals reported relatively more resources for supporting care than did those of either the SNFs or home health agencies. More resources were reported for agencies in high incidence areas.

In order to identify factors inhibiting provision of care to persons with AIDS, a list of nine items was presented. The four most frequently cited are shown (Table 2). Proportionately more CNEs of SNFs indicated that these factors were relevant in their situations.

In response to forced choice options to describe the impact of the epidemic, at

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From the School of Nursing, Oregon Health Sciences University. Address reprint requests to Caroline McCoy White, DrPH, School of Nursing, CB-SN, Oregon Health Sciences University, 3181 SW Sam Jackson Park Road, Portland, OR 97201-2098. This paper, submitted to the Journal May 2, 1990, was revised and accepted for publication October 24, 1990.

TABLE 1—Numbers of Chief Nursing Executives Reporting Agency Experience Caring for Persons with AIDS by Agency Type and Location in County of High or Low Prevalence of AIDS, Oregon, 1989

	Hospitals		SNF		HHA	
Prevalence[a]	High (n = 35)	Low (n = 42)	High (n = 30)	Low (n = 20)	High (n = 16)	Low (n = 36)
Experience with care[b]						
Yes	31	27	3	1	13	16
No	2	8	18	14	2	15
Don't know	2	5	7	4	1	3

[a] On basis of cumulative number of reported cases by county of residence at onset of symptoms (low = 10 or less; high = 11 or more) (Ref. 2).

[b] For hospitals, experience in providing care includes both inpatient and outpatient care. For skilled nursing facilities (SNF) experience refers to inpatient care only. For home health agencies (HHA) experience includes both home care and hospice care.

TABLE 2—Numbers of Chief Nursing Executives Reporting Resources for Care, and Current Issues by Agency Type and Location in County of High or Low Prevalence of AIDS, Oregon, 1989

	Hospitals		SNF		HHA	
Prevalence[a]	High (n = 35)	Low (n = 42)	High (n = 30)	Low (n = 20)	High (n = 16)	Low (n = 36)
Resources for care						
Good compliance with universal precautions	25	29	11	4	11	21
Committee	27	26	10	7	7	11
AIDS specialist	11	8	2	1	3	6
Adequate policies & resources	27	34	10	3	12	24
Current Issues Inhibiting Care of AIDS Patients						
Confidentiality	16	22	13	13	2	11
Inadequate reimbursement	3	10	11	11	5	6
Lack of resources for staff education	2	5	5	3	2	8
Staff, patients & families might shun agency	0	3	14	15	1	5

[a] On basis of cumulative number of reported cases by county of residence at onset of symptoms (low = 10 or less; high = 11 or more) (Ref. 2).

least 40 percent of CNEs in each agency group reported that the epidemic had little or no effect on their agencies. Comments reflected the basis for judgement about low impact: either the agency had not provided care to persons with AIDS or their numbers had been few. However, other comments noted the extensive effort required for writing and implementing policies and procedures for universal precautions.

In assessing anticipated future roles of agencies in response to the epidemic, CNEs of 29 percent of hospitals, 18 percent of SNFs and 37 percent of HHAs anticipated that their agencies would be "at the forefront." This perception was reported by proportionately more CNEs from high prevalence areas where anticipation of being in the forefront was reported by 35 percent from hospitals, 27 percent from SNFs and 46 percent from HHAs.

Discussion

If the Oregon experience is generalizable, agencies in low incidence areas in moderate and low incidence states should expect to provide care for persons with AIDS. And, if SNFs are to be future resources for care for persons with AIDS, the responses of CNEs in these facilities suggest that a concerted effort is needed to market the idea to their sponsors; the fears they expressed point to the need for further community education and involvement.

These data should alert policy-makers about preparedness for providing nursing services to persons with AIDS. At all levels policy-makers must develop a new spectrum of coordinated services across agencies that provides support to nursing care with "models that are effective and attractive to both patients and nurses."[3] ☐

Acknowledgments

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Mass Flea Outbreak at a Child Care Facility: Case Report

Larry D. Corpus, PhD, and Kathleen M. Corpus, PhD

ABSTRACT

Cat fleas, *Ctenocephalides felis* Bouchet, were collected at a Mississippi child care facility after reports of large numbers of adult fleas occurring on children and personnel. One building yielded 161 (99 percent) of the fleas collected. Urticarial lesions due to flea bites occurred on the legs of six children. Flea presence was due to cats occupying the crawl space. Fleas were eradicated by eliminating entry of cats and using residual insecticides throughout the facility. (*Am J Public Health* 1991;81:497–498)

Introduction

Cat fleas, *Ctenocephalides felis* Bouchet, are common obligatory ectoparasites, often evident in homes where pets are kept. Their role in host annoyance and disease transmission is well documented.[1] Mass emergence of adults can occur if population numbers are suitably high and environmental conditions are optimal.[1,2] Outbreaks are rarely noted in the literature, although several entomologists consulted could recall anecdotal occurrences of mass flea emergence. In this paper, we report on a case which occurred at a child care facility in northeast Mississippi.

Methods and Results

Case Report

The child care facility occupies two buildings separated by a playground. Children 6 months to about 3 years of age occupy Building 1; children 3 to 4 years of age use Building 2. Both buildings date from the 1940s and were originally private homes constructed of wood clapboard set upon brick walls surrounding conventional foundations. Plumbing and air handling systems are in crawl spaces with access through doors below exterior stairways.

On July 25, 1988, several children from Building 2 entered the playground. One child returned almost immediately screaming that fire ants were crawling on her legs and biting her. Fire ant mounds had been removed from the playground on several occasions. About 50 insects were estimated to be on the child's legs, and were believed at first to be fire ants. However, when the director brushed the insects off and noticed that they "hopped" onto her arms and that similar insects were on her own pant legs, she concluded that they were fleas rather than fire ants. At the same time, four other children ran to the stairs from the playground indicating that they too, had "fire ants" and "bugs" on their legs and were being bitten. As many fleas as possible were removed from the children and everyone went indoors. A pest control company was contacted, but the premises were not treated until the following day.

On July 27, a maintenance worker entered the crawl space under Building 2 to perform routine equipment maintenance, and was later noted sitting on the outside stairs brushing fleas off his uniform. On July 28, reports were received from parents of children about fleas in their homes "biting everyone."

Physical evidence of flea bites was reported after August 2. Six children showed advanced cases of ulcerative sores on their legs, with the number of bites ranging from one to four per leg. Four of the six children exhibiting flea bites were in the initial group which entered the playground on July 25. Most children were not affected since facility personnel chose to remain indoors from July 25 to August 12, a period of 14 days. No adults reported any lasting effects of flea bites beyond brief itching around their ankles and knees. Only one child from Building 1 exhibited any flea bite symptoms, bearing a group of three urticarial lesions on the inside of his right ankle.

The senior author, notified of the flea problem on August 4, visited the site on August 5. Personnel were interviewed and collection of fleas initiated. While standing in the playground, 35 adults fleas were removed by aspiration from pant legs and socks. It was discovered that only a narrow band around each building had been sprayed with chlorpyrifos by the pest control company on July 26. On August 12, after all children, parents, and personnel had left the premises for the weekend, the entire facility was treated with propetamophos (Safrotin) and precor indoors, propetamophos and chlorpyrifos outdoors including the playground, and diazinon dust throughout the crawl spaces.

Adult fleas were removed from clothing and placed into 70 percent ethanol. Litter samples were taken from under each building and throughout the playground and berlesed (heat extraction) for

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Address reprint requests to Larry D. Corpus, PhD, Department of Entomology, Drawer EM, Mississippi State University, Mississippi State, MS 39762. Dr. K. Corpus is with the Department of Home Economics at MSU. This paper, submitted to the Journal January 22, 1990, was revised and accepted for publication September 5, 1990.

TABLE 1—Numbers of Adult and Larval (in parentheses) Fleas Collected from Building 2

	Collecting Method		
Date	Berlese	Dry Ice	Aspiration
Aug 5	-	-	35
Aug 8	5 (5)	2	43
Aug 10	17 (8)	4	21
Aug 15	8 (2)	1	7
Aug 17	0 (1)	0	2
Aug 19	0	0	0
Aug 22	0	0	0
Aug 24	0	0	0
Aug 26	0	0	0
Total	30 (16)	7	108

two weeks. Dry ice in white enamel pans was placed in the crawl spaces below each building. Continued collection and monitoring of fleas occurred until August 26 even though no fleas were detected in any stage after August 16.

Identification of the fleas indicated that they were cat fleas, *Ctenocephalides felis* Bouchet (Siphonaptera:Pulicidae). The number of fleas collected from Building 2 for the period August 5 to August 26 are shown in Table 1. Adult fleas comprised the majority of individuals collected (90.2 percent), whereas larvae, all from the crawl space, accounted for only 9.8 percent of the total. Only two adult fleas were collected adjacent to Building 1. The disproportionate number of flea larvae and adults collected from Building 2 suggests that adult fleas were originating from below Building 2. Both the maintenance worker and the senior author attracted fleas while under Building 2, but not while under Building 1. Most of the fleas (98.6 percent) were collected during the first four collecting dates. No flea pupae were collected.

Discussion

The most probable explanation for the greater number of fleas around Building 2 was the presence of several cats in the vicinity on several previous occasions and their continued use of the crawl space. Cats from surrounding buildings roam freely during the day and night and had been trapped near Building 2 on several previous occasions. Fur, feces, and masticated rodent remains below Building 2 further attested to the presence of cats. Cats had entered through open vent holes in the foundation and occupied the crawl space so that eventually a substantial population of fleas had built up. Since similar conditions did not exist at Building 1, very few fleas occurred near it.

Although *Ctenocephalides* spp. fleas have the potential to vector a number of diseases,[2,3] those collected in this case were not tested for pathogens. It was recommended that, in the future, better chemical control measures be used in a more timely manner and that all openings into the crawl spaces below each building be closed off with at least a fine mesh screen. We further suggested that trapping of free-ranging pets be continued, and leash ordinances be enforced in order to reduce or eliminate potential flea hosts. □

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Any mention of a proprietary product does not constitute an endorsement.

Mississippi Agricultural and Forestry Experiment Station Contribution Number J-7325.

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Impact of Work on the Quality of Life of Community-Residing Young Elderly

Hirsch S. Ruchlin, PhD, and John N. Morris, PhD

ABSTRACT

Using a random sample of 310 Massachusetts community-residing elderly between the ages of 65 and 74, this study investigates the relationship between employment status and quality of life using a modified version of the Spitzer Uniscale QL index. The odds of reporting the highest quality of life rating, after controlling for socioeconomic and health characteristics, was 3.51:1 for those who worked versus those who did not do so. (*Am J Public Health* 1991;81:498–500)

Introduction

A desire to work beyond the expected age of retirement has become more prominent in recent years.[1,2] Predictions of labor shortages and growing competition for those willing and able to work[3] has directed attention to the continued labor force participation of older people and to the potential reemployment of retirees. Maloney and Paul assert that greater opportunities to work could improve the quality of life for older people.[4] They note that nearly half of the nonworking adults between the ages of 55 and 64 reported that they would have preferred to continue working. Those who stopped work involuntarily said that they were less sat-

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Address reprint requests to Hirsch S. Ruchlin, PhD, Professor of Economics in Public Health, Department of Public Health, Cornell University Medical College, 1300 York Avenue, Box 4, New York, NY 10021. Dr. Morris is with the Hebrew Rehabilitation Center for Aged, Boston. This paper, submitted to the Journal May 2, 1990, was revised and accepted for publication August 6, 1990.

TABLE 1—Distribution of Quality of Life Ratings

Quality of Life Rating	Total	Working	Not Working
1 (lowest)	0	0	0
2 (low)	1	0	1
3 (low middle)	0	0	0
4 (middle)	17	1	16
5 (high middle)	14	2	12
6 (high)	79	8	71
7 (highest)	199	66	133
Total	310	77	233

TABLE 2—Factors Associated with the Odds of Reporting the Highest Quality of Life Rating

Variables	Odds Ratio (95% CI)
Sex	1.02 (0.57,1.81)
Marital Status	1.41 (0.79,2.53)
Self-Assessed Health Status	3.39 (1.78,6.48)
Functional Status	8.29 (2.18,31.48)
Self-Assessed Financial Status	1.55 (0.88,2.74)
Work Status	3.51 (1.67,7.34)

CI = Confidence Interval

isfied with their lives than those who chose either to continue work or to retire.

Evidence to support the hypothesis that work has a positive impact on the elderly's quality of life is limited. Most of the literature focuses on whether retirement leads to increased mental health problems.[5-12] One study conducted by Soumerai and Avorn adopted an employment rather than a retirement focus.[13] Based on a sample of 55 individuals (25 experimentals and 30 controls) they report higher life satisfaction index scores for the experimental (i. e. work) group.

We report here a study of the relationship between employment and quality of life among those elderly who are most likely to be interested in working—those in the age group 65 to 74.

Methods

The Sample

Our data come from a representative sample of Massachusetts community elderly assembled by the Department of Social Gerontological Research, Hebrew Rehabilitation Center for Aged which spans the period 1982–87 (baseline n = 2,682). The sample was constructed by: dividing the state into its 27 home care areas; stratifying areas by size (small, medium, and large) and randomly selecting cities and towns from each area; and randomly selecting for each selected area residents 60 years of age and older from within two age strata (under 70 and 70 and over) from the state census. Data in the sample were obtained through telephone interviews.

The analysis in this study is limited to 1987 data and to individuals under age 75 in 1987 (n = 351). Individuals not in a community setting (n = 24) or who never worked (n = 17) were excluded, leaving a study sample of 310 people: 77 employed (24.8 percent) and 233 (75.2 percent) not employed. The mean age of these two groups was 69.8 years (SD ± 2.3) and 70.3 years (SD ± 2.3). Among those employed, 21 (27.3 percent) worked full-time and 56 (72.7 percent) part-time.

Study Variables

Quality of life, the study's dependent variable, was measured by a self-reported, broad-based, well-established measure—the Spitzer Uniscale QL index—using the scoring metric developed by Morris and Spitzer for use in a general community population. The Uniscale QL has been used with a number of different elderly populations including those with and without serious disabilities,[14,15] and has high interrater reliability and excellent validity properties when compared against other quality of life indicators.[14] It is described in Appendix A.

The independent variables used were: sex, marital status, a self-assessed overall health rating, functional status, a self-assessed rating of financial status, and work status. The coding of these variables and their descriptive statistics are reported in Appendix Table A2.

Results

Quality of Life Rating by Work Status

The distribution of quality of life ratings are reported in Table 1. Approximately 90 percent of the sample reported quality of life ratings in one of the two top categories. By work status, 85.7 percent of the working elders (95% CI = 77.9, 93.5) reported the highest quality of life rating compared to 57.1 percent of the nonworking elders (95% CI = 50.7, 63.5). Focusing on the two highest quality of life ratings, 96.1 percent of the working elders (95% CI = 91.8, 100) were in this range compared to 87.6 percent (95% CI = 83.4, 91.8) of the nonworking elders. No significant differences emerged in a separate analysis of quality of life ratings by extent of work (i.e. full- or part-time).

Impact of Work on Quality of Life

Logistic regression was used to assess the impact of the independent variables on the likelihood of reporting the highest quality of life rating of 7 versus ratings of 1 through 6. The odds ratios associated with each of these variables are reported in Table 2. Three variables display definite impacts—self-assessed health, functional status, and work status. Adjusting for other factors included in the analysis, the likelihood of an elderly person who is working either full-time or part-time displaying the highest quality of life rating is 3.51 times that of an elderly person who is not working.

Discussion

Two health-related factors emerged as associated with the highest quality of life profile: self-assessed health, and functional independence. Working also emerged as a very important element. While work produces extra income to supplement Social Security and pension or investment income, financial security per se was not a significant correlate of a very high quality of life profile. We suggest that the feeling of "usefulness" and "value" traditionally associated with having a job underlies the work-quality of life relationship reported here.

Our results, which clearly await replication from studies that span a broader geographic base and time horizon, indicate that efforts at improving the life of the growing number of "young elderly" (i. e. people in the 65–74 age range) encompass programs that help people over the age of 65 stay on their current job or find new employment that better meets their current needs and capabilities. Such efforts will contribute to the improved social and mental health of the elderly. ☐

Acknowledgment

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APPENDIX

In its self-report version, Uniscale QL scores quality along a seven-category continuum based on a response to the following question: "How would you rate your overall quality of life? Think of a ladder with 7 rungs, where the first, lowest rung applies to someone completely dependent physically on others, seriously impaired mentally, unaware of surroundings, and in a hopeless position. At the top of the ladder, the 7th rung applies to someone physically and mentally independent, communicating well with others, able to do most of the things enjoyed, pulling one'sown weight, with a hopeful yet realistic attitude. Where would you place yourself on this 7-rung ladder?" The distribution of the self-report Uniscale QL index for a representative sample of community elders in Massachusetts is reported in Table A1.

TABLE A1—Uniscale QL Distribution for A Representative Sample Of Community Elders in Massachusetts (N = 1,582) (in percent)

		Extent of Functional Disability		
QL Rating	Total Sample	No ADL/IADL Deficit	IADL Deficit/ No ADL Deficit	ADL Deficit
1 (lowest)	0.8	0.1	0.3	11.2
2 (low)	0.6	0.0	0.5	8.8
3 (low middle)	1.3	0.1	5.1	10.5
4 (middle)	4.6	3.7	7.0	12.0
5 (high middle)	20.1	16.3	40.5	33.4
6 (high)	32.2	33.9	31.9	8.7
7 (highest)	40.4	45.9	14.7	15.4
(% of sample)		(82.6)	(11.7)	(5.7)

TABLE A2—Coding of Independent Variables and their Descriptive Statistics

Variable	Coding	Mean	Standard Deviation
Sex	1 = Female 0 = Male	0.558	0.497
Marital Status	1 = Married 0 = Not married	0.600	0.491
Self-Assessed Health Status	1 = Excellent or good 0 = Fair or poor	0.781	0.414
Functional Status	1 = Independent 0 = Dependent[a]	0.919	0.273
Self-Assessed Financial Status[b]	1 = Good 0 = Fair or poor	0.687	0.464
Work Status	1 = Working 0 = Not working	0.248	0.433

N = 310 except for functional status where n = 309.

a) Dependent in either personal activities of daily living and/or instrumental activities of daily living.

b) Data for this variable were drawn from the responses to the following question: "Would you say that you have enough money to live on with little trouble, just enough to get by, or not enough to make ends meet?" These responses were characterized as good, fair or poor, respectively.

Black-White Differences in Cancer Prevention Knowledge and Behavior

Christopher Jepson, PhD, Larry G. Kessler, ScD, Barry Portnoy, PhD, and Tyson Gibbs, PhD

ABSTRACT

Data from the 1987 National Health Interview Survey Cancer Control Supplement were used to estimate multivariate logistic regression models of diet change, mammography utilization, stool blood test utilization, and smoking. Predictor variables included race, sex, age, income, dietary concerns, and four knowledge-related variables: education and three measures of cancer prevention knowledge. When knowledge variables were included in the models, race was not a significant predictor of behavior, with one exception: among women, Blacks were found to smoke less than Whites. (*Am J Public Health* 1991;81:501–504)

Introduction

Blacks have a higher incidence of and mortality from cancer than Whites.[1,2] This differential presumably reflects differences in demographic factors such as education and income, and cognitive psychological factors such as attitudes and knowledge which influence behavior.

Although studies have shown that Blacks tend to have lower levels of cancer knowledge than Whites,[3–5] the relationship between knowledge level and cancer-preventive behavior has not been adequately documented. Some studies indicate a relationship between knowledge and behavior,[6–9] while others do not do so.[10,11] At minimum, however, some basic preexisting level of knowledge is a prerequisite for any instrumental behavior. Theories such as the Health Belief Model,[12,13] PRECEDE,[14] and Fishbein and Ajzen's theory of reasoned action[15] also postulate a relationship between behavior and beliefs, which are presumably based upon knowledge to some extent.

In 1987 the National Health Interview Survey (NHIS)[16] included a Cancer Control Supplement assessing a variety of demographic and cognitive psychological factors in addition to cancer-preventive behaviors. This made it possible to examine the extent to which such factors were associated with Black-White behavioral differences.

Methods

A total of 22,043 adult respondents over the age of 18 provided data on the 1987 NHIS Cancer Control Supplement. To increase accuracy of point estimates, Blacks were oversampled: 3,053 respondents (13.9 percent of the total sample) described themselves as Black.

For the purposes of our analyses, the study population was restricted to respondents describing themselves as either Black or White; Hispanics and members of other ethnic or racial groups were excluded. This reduced the total number of respondents to 20,088. In the analyses reported here, the data are weighted (using weights assigned by the National Center for Health Statistics) so as to be representative of the US population.[16]

We selected four measures of cancer-preventive behavior as dependent variables. These variables are described in Table 1.

The independent variables selected for our analysis fell into two main categories—demographic (race, sex, income, education, age) and cognitive psychological (knowledge, belief, and attitude) measures. They are described in Table 2. A detailed description is provided in the Appendix.*

In our data analysis, we first examined Black-White differences in the dependent and independent variables, and then estimated logistic regression models of each dependent variable, using a special software program, RTILOGIT,[17] designed specifically for use on data sets of this type.

Results

Black-White Differences

Table 1 reveals substantial Black-White differences in rates of diet change, mammography, and stool blood testing; smoking differences are essentially confined to men. Table 2 shows that Blacks scored lower than Whites on the three knowledge measures and were less likely to believe that cancer is preventable.

Some independent variables, such as the knowledge measures, are continuous. Mean Black-White differences on these variables, shown in Table 2, appear small but when they are dichotomized and percentages are compared the differences are

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*Additional information on standard errors of point estimates presented in this paper, odds ratios, etc., are available upon request from the first author.

From the National Cancer Institute, Bethesda, MD. Address reprint requests to Christopher Jepson, PhD, University of Pennsylvania School of Nursing, Room 309, Nursing Education Bldg., Philadelphia, PA 19104. This paper, submitted to the Journal November 8, 1989, was revised and accepted for publication August 13, 1990.

TABLE 1—Distribution of Dependent Variables by Race[f]

Variable	Description	Race[a]		OR[b] (95% CI)
		Black (n)	White (n)	
Diet change	Self-report of whether respondent has "ever made any lasting and major changes in [diet] for health reasons"	31.3 (2971)	36.8 (17089)	0.78 (0.70, 0.87)
Mammography	Self-report of whether respondent has ever had a mammogram[c,d]	31.2 (907)	40.0 (5524)	0.68 (0.57, 0.81)
Stool blood test	Self-report of whether respondent has ever had a stool blood test[c,e]	31.3 (1448)	39.4 (9320)	0.70 (0.60, 0.82)
Smoking status	Self-report of whether respondent is currently a cigarette smoker	32.4 (2868)	28.7 (16521)	—
	Men only	37.8 (1020)	30.2 (7025)	1.40 (1.15, 1.71)
	Women only	27.9 (1848)	27.2 (9496)	1.04 (0.89, 1.20)

[a] Figures are percentages of Blacks and Whites reporting the given behavior. Respondents who describe themselves as Hispanic or as neither Black nor White are excluded.

[b] Odds ratio = odds of Blacks (vs Whites) reporting the given behavior.

[c] Respondents were first asked if they had heard of the given test. Those answering "no" or "don't know" were counted as never having had the test.

[d] Only women age 40 and over responded.

[e] Only men and women age 40 and over responded.

[f] Data are weighted to be representative of US population.

equal to or greater than those seen for the categorical variables.

Regression Analyses

In order to determine the extent to which knowledge-related factors could account for the relationships between race and cancer-preventive behaviors, we estimated two multivariate logistic regression models of each dependent variable. In the first or reduced model, all independent variables are included except for the four knowledge-related variables (education and the three measures of cancer prevention knowledge). The second or full model includes all of the independent variables in Table 2.

Table 3 shows the coefficients (betas) for all independent variables in the full models of the four dependent variables. The coefficients for race translate into the following odds ratios (i.e. the odds of Blacks reporting the given behavior compared to Whites): for diet change, 0.94 (95% CI = 0.84, 1.05); for mammography, 0.94 (95% CI = 0.77, 1.16); for stool blood test, 1.08 (95% CI = 0.91, 1.28); for smoking among men, 1.03 (95% CI = 0.83, 1.29); and for smoking among women, 0.71 (95% CI = 0.59, 0.85). Thus, in all models except that of women's smoking, Black-White differences in behavior do not persist when the other predictors are held constant.

Among women, race is not a significant predictor of smoking status by itself, as shown in Table 1. As Table 3 shows, however, Black-White differences become significant in the full model.

TABLE 2—Distribution of Independent Variables by Race[b]

Variables	Description	Percent in Population or mean value[a]		
		Overall	Black	White
Demographic Factors				
Race	% White	88.1	—	—
	% Black	11.9	—	—
Sex	% Male		44.9	47.7
	% Female		55.1	52.3
Education	Years completed, from 0 (none) to 18 (18 or more)		11.4	12.7
Age	Age in years		40.8	44.6
Square of age			—	—
Family income:				
% Under $10,000			27.1	11.2
% $10,000–$19,999			22.4	18.8
% $20,000–$34,999			20.8	27.3
% $35,000 & over			14.8	31.5
% Unknown			11.9	11.2
Cognitive Psychological Factors				
Fat & fiber knowledge	Maximum score = 8		3.65	4.79
Risk factor knowledge	Maximum score = 7		3.63	4.46
Cigarette risk knowledge	Maximum score = 5		3.62	4.12
Dietary concerns	Maximum score = 3		2.16	1.95
Belief that cancer is *not* preventable:				
% Agree/strongly agree			30.8	20.5
% Disagree			45.0	53.7
% Strongly disagree			9.5	16.9
% No opinion			14.8	9.0

[a] Respondents who describe themselves as Hispanic or as neither Black nor White are excluded. When percentages are given, the variable in question is a dummy variable; when means are given, the variable is continuous.

[b] Data are weighted to be representative of US population.

In the reduced models of each dependent variable, the betas for race are intermediate between those shown in Table 1 and those in Table 3, and all are significantly different from zero.

Discussion

Our results were generally consistent across all dependent variables except for smoking status among women. By includ-

TABLE 3—Beta Values for Independent Variables in Models of Cancer-Preventive Behavior[a]

Variables	Diet change (n = 19596)	Mammography utilization (n = 6238)	Stool blood test utilization (n = 10438)	Smoking status	
				Men (n = 7885)	Women (n = 11333)
Race	−0.07	−0.06	0.07	0.03	−0.35***
	(−0.18, 0.05)	(−0.26, 0.15)	(−0.10, 0.24)	(−0.19, 0.26)	(−0.52, −0.17)
Sex	0.05	—	−0.04	—	—
	(−0.03, 0.12)		(−0.14, 0.05)		
Age	0.05***	0.13***	0.19***	0.13***	0.12***
	(0.04, 0.06)	(0.08, 0.18)	(0.15, 0.22)	(0.10, 0.15)	b
Square of age	−0.0004***	−0.001***	−0.001***	−0.002***	−0.002***
	b	b	b	b	b
Family income:					
Under $10,000	0.06	−0.32***	−0.04	0.24*	0.49***
	(−0.07, 0.20)	(−0.50, −0.13)	(−0.21, 0.13)	(−0.01, 0.48)	(0.28, 0.71)
$10,000–$19,999	−0.06	−0.16	0.03	0.28**	0.33***
	(−0.18, 0.07)	(−0.36, 0.04)	(−0.11, 0.16)	(0.04, 0.52)	(0.15, 0.51)
$20,000–$34,999	−0.05	0.15	0.05	0.09	0.20*
	(−0.17, 0.08)	(−0.08, 0.38)	(−0.10, 0.20)	(−0.13, 0.31)	(−0.00, 0.39)
$35,000 & over	−0.003	0.34***	0.18**	−0.13	−0.05
	(−0.13, 0.12)	(0.12, 0.56)	(0.04, 0.31)	(−0.37, 0.11)	(−0.22, 0.12)
Dietary concerns	0.52***	0.07**	0.07***	−0.23***	−0.17***
	(0.48, 0.56)	(0.00, 0.13)	(0.02, 0.11)	(−0.29, −0.18)	(−0.22, −0.12)
Belief that cancer is *not* preventable:					
Agree/strongly agree	−0.15**	0.26**	−0.03	0.21	0.31***
	(−0.29, −0.01)	(0.04, 0.49)	(−0.19, 0.13)	(−0.04, 0.47)	(0.11, 0.50)
Disagree	0.12*	0.23**	0.08	0.01	0.11
	(−0.01, 0.24)	(0.02, 0.45)	(−0.09, 0.25)	(−0.22, 0.24)	(−0.10, 0.32)
Strongly disagree	0.53***	0.23*	0.09	−0.06	0.09
	(0.39, 0.68)	(−0.01, 0.48)	(−0.12, 0.30)	(−0.33, 0.21)	(−0.15, 0.33)
Education	0.02***	0.07***	0.06***	−0.11***	−0.10***
	(0.00, 0.03)	(0.04, 0.09)	b	(−0.13, −0.09)	(−0.12, −0.08)
Fat & fiber knowledge	0.09***	0.10***	0.11***	−0.01	−0.02
	(0.07, 0.12)	(0.06, 0.14)	(0.07, 0.14)	(−0.05, 0.02)	(−0.05, 0.02)
Risk factor knowledge	0.08***	0.01	0.02	−0.03	−0.01
	(0.05, 0.10)	(−0.02, 0.05)	(−0.01, 0.05)	(−0.07, 0.01)	(−0.05, 0.02)
Cigarette risk knowledge	0.04**	0.09***	0.09	−0.15***	−0.19***
	(0.01, 0.06)	(0.04, 0.13)	(0.05, 0.13)	(−0.20, −0.11)	(−0.24, −0.15)

[a] Figures are regression coefficients (betas) and 95% confidence intervals (in parentheses).

[b] 95% confidence limits are within .01 of parameter estimate.

*$p < .10$

**$p < .05$

***$p < .01$

ing knowledge-related variables in the analysis, Black-White behavioral differences were eliminated. In the case of women's smoking status, when other factors are controlled for, Black women become less likely than White women to smoke.

In addition to accounting for the effects of race, knowledge-related variables also displayed independent relationships with cancer-preventive behaviors. Education was a significant predictor in all models. Cigarette risk knowledge also contributed significantly in all cases, with its contribution greatest in the models of smoking status. Fat and fiber knowledge was a significant predictor of all dependent variables except smoking status; somewhat surprisingly, it was no more significant for diet change than for the other dependent variables. Risk factor knowledge contributed significantly only to the model of diet change.

The cross-sectional nature of the analysis places significant limitations on the conclusions to be drawn from this study. Associations discovered in such a study cannot be assumed to reflect causal relationships, and it is possible that other variables could account for Black-White behavioral differences as well as or better than the ones we chose to examine. Nevertheless, this study is unique in that it is the first national study assessing the independent effects of knowledge, beliefs, attitudes, and demographic factors on Black-White differences in cancer-preventive behaviors. The importance of cognitive psychological factors is indicated by the fact that even the imperfect measures available to us eliminated the behavioral differences.

Our findings suggest that a fuller understanding of Black-White behavioral differences can be gained by including measures of knowledge and attitudes, especially if these measures assess specific concepts directly related to the behaviors in question. Second, they suggest that there may be merit in investigating the use of interventions designed to narrow the behavioral discrepancies between Blacks and Whites by reducing the knowledge discrepancies. Longitudinal studies of this type should answer the question of whether enhancing cancer knowledge will result in increased cancer-preventive behaviors as a stepping stone to reductions in cancer mortality. ☐

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APPENDIX

Description of Independent Variables

Reported family income was treated as a categorical variable rather than as an ordinal variable because approximately 12 percent of respondents did not provide income data. Respondents who did provide income data were divided into four categories, and each group was assigned a dummy variable; the unknowns served as the reference category.

Three measures of cancer prevention knowledge were created as follows: The first, *fat and fiber knowledge,* represents the number of correct answers to two questions in which respondents were given lists of foods and asked to identify which were sources of fat and fiber. The second, *risk factor knowledge,* was similarly created from an item in which respondents were asked to identify cancer risk factors. The third, *cigarette risk knowledge,* was taken from a set of five items in which respondents were asked to identify diseases caused by cigarette smoking. The reliability of this measure was very high (Cronbach's alpha = 0.90).

An attitude measure of dietary concerns was created from a set of three items in which respondents were asked if avoiding sodium, avoiding sugar, and lowering cholesterol were important concerns for them (alpha = 0.80).

Belief in the preventability of cancer was measured by degree of agreement with the statement, "There is very little a person can do to reduce his or her chances of getting cancer." Respondents were divided into four categories—those who agreed or strongly agreed, those who disagreed, those who strongly disagreed, and those with no opinion—with the last serving as the reference category.

Report of an Association between Race and Thyroid Stimulating Hormone Level

Joel M. Schectman, MD, MPH, Gene A. Kallenberg, MD, Robert P. Hirsch, PhD, and Ronald J. Shumacher, MD

ABSTRACT

We examined the association between race and thyroid stimulating hormone (TSH) level in 809 consecutive Black or White adults tested for mild suspicion of hypothyroidism with a normal TSH result. The mean TSH level of Blacks was 0.4 (SE .053) mU/L lower than that for Whites after age and sex adjustment, race explaining 6.5 percent of the variation in TSH levels. A validation sample yield similar results. This finding supports the possibility that differences in thyroid function and/or regulation may be associated with race. (*Am J Public Health* 1991;81:505–506)

Introduction

In two recent reports the incidental finding of an association between race and primary hypothyroidism has been noted.[1,2] The odds of both borderline and substantial thyroid stimulating hormone (TSH) elevations were greater among Whites than Blacks, controlling for age and sex. An increased risk of autoimmune thyroiditis among Whites has also been mentioned in thyroid textbooks,[3,4] but supported mainly by old hospital prevalence studies[5,6] with biases likely favoring detection of disease among Whites. Age and sex associations for primary hypothyroidism and autoimmune thyroiditis are well accepted, with previous studies documenting corresponding age and sex associations with TSH levels as well.[7-9] We hypothesized that a similar association would be found between race and TSH levels if there were a substantial racial association for hypothyroidism and autoimmune thyroiditis.

Methods

Subjects were consecutive adult primary care patients of the George Washington University Health Plan that had a TSH level for suspicion of hypothyroidism due to nonspecific symptoms or signs from April 1985 to March 1987. Patients on medications known to affect thyroid function tests or with a history of thyroid disease were excluded. Of 880 such patients identified, 33 were excluded due to an abnormal TSH level and 38 because either no race or one other than Black or White was noted in the medical record. This analysis is restricted to the remaining 809 Black or White patients tested for suspicion of hypothyroidism and found to have TSH levels within normal limits. Categorization of race was generally by patient self-report on the data base form in the chart. The racial demographics of this sample were similar to those of the entire health plan patient population.

The Magic TSH RIA (CIBA Corning Magnetic Immunochemistries, Medfield, MA) with normal limits of 0–5 mU/L was utilized for all TSH determinations obtained during the initial study period. A linear regression analysis was used to test for an association between race and TSH level. The model utilized TSH as the dependent variable, with age, sex, and race as independent variables, and chronic medical problems and/or medication use (both treated as dichotomous variables) as potential confounding factors. An analysis utilizing a log transformation of TSH to normalize the skewed distribution yielded similar results. Interactions between the independent variables were evaluated and a P-value of 0.10 or less was required to remain in the final model in a stepwise procedure.

In August 1989, a validation sample of 97 patients with normal TSH levels was collected at the study site by the same methods. The sample size was calculated to assure a 90 percent power of detecting a 0.4 mU/L difference in mean TSH levels at a 0.05 level of significance. The Magic mab TSH (CIBA Corning Magnetic Immunochemistries), an immunoradiometric sandwich assay with normal limits of 0.3–4.5 mU/L, was utilized for the TSH determinations on this patient sample. The analysis proceeded as described above.

Results

Of the patients from the initial study sample, 406 were Black and 403 were White. The interaction terms did not achieve a P-value of 0.10 or less and were thus eliminated from the final model. Chronic medical conditions and medica-

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From the George Washington University Medical Center (Schectman, Kallenberg, Hirsch) and the Georgetown University Hospital (Shumacher). Address reprint requests to Joel M. Schectman, MD, MPH, Department of Health Care Sciences, George Washington University Medical Center, 2150 Pennsylvania Avenue, NW, Rm 2B-420, Washington, DC 20037. This paper, submitted to the Journal March 8, 1990, was revised and accepted for publication August 6, 1990.

TABLE 1—Linear Regression Model of TSH as a Function of Race

Variables	Sample	Coefficient	Standard Error	P-value
Age*	Initial	0.007	0.002	0.004
	Validation	0.006	0.006	0.32
Sex**	Initial	−0.081	0.074	0.27
	Validation	−0.45	0.23	0.06
Race***	Initial	0.402	0.053	<0.0001
	Validation	0.45	0.18	0.01

*In years.
**Coded: female = 0, male = 1.
***Coded: Black = 0, White = 1.

tion use were not significantly associated with TSH levels. Age and race, but not sex, were independently predictive of TSH levels within the normal range (Table 1). Race explained 6.5 percent and age 0.8 percent of the variation in TSH levels. Whites had an age- and sex-adjusted mean TSH level 0.4 mU/L higher than Blacks.

Of the 97 patients in the validation sample, 51 were Black and 46 were White. The regression analysis again reveals an association between TSH level and race independent of sex and age. Race explained 5.5 percent of the variation in TSH levels with Whites having an age- and sex-adjusted mean TSH level 0.45 mU/L higher than Blacks.

Discussion

The present analysis reveals a racial association for TSH levels within the normal range in a population of primary care patients tested for generally mild suspicion of thyroid disease. The fairly large mean difference in TSH levels between Blacks and Whites in the two samples suggests that this association may be of clinical importance and is consistent with previously reported findings of a higher prevalence of borderline and frankly elevated TSH levels in Whites.[1,2] The association between race and TSH level was much stronger than that for age or sex, factors known to be associated with TSH level. An older sample would likely have demonstrated a greater age and sex dependency, however.

It is unlikely that these results are due to chance, given the highly significant P-value and the similar results obtained in the validation sample using the newer super-sensitive TSH assay.

It is possible that biased selection could account for these results since neither data set resulted from a random sampling of the population. This was necessary since, to our knowledge, no data base is available that contains race and TSH data for a normal reference population. However, for such a selection bias to have occurred, presumably euthyroid Whites with TSH levels higher in the normal range than Blacks (controlling for age and sex) must have been preferentially tested in both study samples. We are unaware of any reason for this to have occurred.

Several possible explanations of a racial association for normal TSH levels can be suggested. Environmental factors such as iodide consumption[4] might vary sufficiently between Blacks and Whites in the study population to affect thyroid function and account for the racial association. It is also possible that genetic factors differ between the Black and White patients tested and could account for a difference in TSH levels. Given previously determined relationships between race, gene frequency, and autoimmune disease,[10,11] such an association would not be surprising and has in fact been suggested by one study of post-partum thyroid dysfunction.[12] Population-based data are needed to confirm an association between race and thyroid function, as well as establish a link with antithyroid antibodies or other factors. Such information should be available from the NHANES III database in the future. ☐

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Association of Waist to Hip Ratio and Family History with the Prevalence of NIDDM among 25,272 Adult, White Females

Robert D. Morris, PhD, and Alfred A. Rimm, PhD

ABSTRACT

The independent association of waist to hip ratio (WHR), relative weight, and family history with the prevalence of non-insulin dependent diabetes mellitus (NIDDM) was investigated in a population of 25,254 White females in the United States and Canada. A multivariate logistic analysis yielded standardized odds ratios of 1.47, 1.46, 1.54, and 1.14 for family history index, WHR, relative weight, and age, respectively. (*Am J Public Health* 1991;81:507–509)

Introduction

The association between obesity and non-insulin dependent diabetes mellitus (NIDDM) is well established.[1-3] More recent studies have demonstrated an association between the distribution of adipose tissue and NIDDM after adjusting for obesity and other risk factors.[4-11] The studies cited have used the ratio of waist to hip girth (WHR) as a measure of body fat distribution as first proposed by Rimm and Hartz.[4] The present study evaluates the independent associations of WHR, family history, relative weight, and age as they relate to the prevalence of NIDDM.

Methods

In 1969, TOPS (Take Off Pounds Sensibly), an organization of weight conscious women devoted to a responsible approach to weight control, sponsored the administration of an extensive health inventory to their members. The study yielded a total of 73,532 valid completed questionnaires from White females. Further details of the study design and the type of data collected are given elsewhere.[12,13]

Determination of the presence or absence of diabetes was based upon the subjects response of yes or no to the question, "Has a doctor ever said you had diabetes?" A study to assess the validity of the data, indicated that the medical information and anthropometric measurements supplied by the women was not biased by under- or overreporting.[14]

Only those women between the ages of 40 and 70 at the time of the study were included in the present analysis because the majority of cases of NIDDM appear in this age range. The exclusion of insulin dependent women resulted in a total of 25,272 White female subjects.

Weights and current girth measurements were self-reported. Relative weight taken as the ratio of actual weight to ideal body weight for height as given in the Metropolitan Life Insurance Tables[15] medium frame size served as an index of obesity. Relative weight ranged from 80 to 312 percent of ideal weight with a mean value of 1.32 and a median of 1.27. The ratio of waist to hip girth was used as the measure of body fat distribution.

The women recorded presence or absence of a number of diseases in their parents, grandparents, siblings, and children. This was used to calculate a family history index (FHI) for diabetes.[16] The following formula gives the index in terms of presence (= 1) or absence (= 0) of diabetes in the mother (M), the father (F), the brothers (B), the sisters (S), paternal grandparents (PGP), and maternal grandparents (MGP) of the subject.

$$\mathrm{FHI} = (M + F + B + S) + (PGP + MGP)/2.$$

Because of recent evidence associating smoking with increased WHR,[17] it was included in this analysis. Respondents were categorized as nonsmokers, light to moderate smokers, and heavy smokers based upon the number of pack-years smoked.

Subjects were divided into tertiles for WHR (<0.75, 0.75–0.80, and >0.80); they were divided into three groups for both FHI (0.0, 0.5–1.0, and >1.0), and for relative weight (<1.25, 1.25–1.50, and >1.50). This produced a $3 \times 3 \times 3$ matrix with 27 cells. Diabetes prevalence was determined for each of these cells.

Multiple logistic regression was used to evaluate the prevalence of diabetes as a function of age, family history index, WHR, relative weight, and all possible two and three way interactions. While this approach does not delineate interaction to

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Address reprint requests to Alfred A. Rimm, PhD, Division of Biostatistics and Clinical Epidemiology, Medical College of Wisconsin, 8701 Watertown Plank Road, Milwaukee, WI 53226.
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TABLE 1—Comparison of Selected Characteristics between Diabetic and Nondiabetic Women from the 1969 National Survey of Members of TOPS Club, Inc*

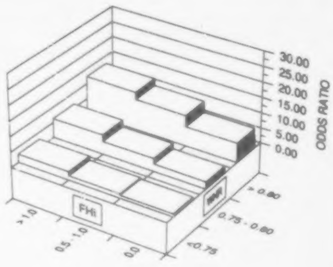
Variables	Diabetic	Nondiabetic	Total population
n	1,345	23,927	25,254
% of total	5.3	94.7	100
Average Age (years)†	51.6 ± 7.4	50.3‡ ± 7.0	50.4 ± 7.1
Average Family History Index††	0.72 ± 0.72**	0.31‡ ± 0.31	0.33 ± 0.33
Average Relative Weight	1.45 ± 0.29	1.31‡ ± 0.25	1.31 ± 0.25
Average Waist: Hip Ratio	0.81 ± 0.07	.78‡ ± 0.06	0.78 ± 0.06

*TOPS Club, Inc., Take Off Pounds Sensibly Club, Inc.
†Mean ± one standard deviation.
‡$p < .001$ for univariate analysis comparing diabetic to nondiabetic women.
††Family history index = no. of diabetic parents and siblings + (no. of diabetic grandparents ÷ 2).
**The distribution of the family history index approximates an exponential density distribution in which case the standard deviation is approximated by the mean.

TABLE 2—Prevalence of NIDDM among Women from the 1969 National Survey of Members of TOPS Club, Inc,* when Stratified by Relative Weight, Family History Index for Diabetes, and Waist to Hip Ratio

Relative Weight	Family History Index	WHR <0.75	WHR 0.75–0.80	WHR >0.80
0.8–1.25	0.0	1.4 (4484)**	2.0 (2711)	3.5 (1768)
	0.5–1.0	3.2 (1250)	5.3 (811)	8.6 (452)
	1.5–5.0	8.2 (352)	8.5 (235)	16.2 (179)
1.26–1.50	0.0	1.9 (1996)	3.0 (2130)	6.5 (1539)
	0.5–1.0	5.5 (550)	6.8 (759)	11.6 (934)
	1.5–5.0	12.3 (199)	9.1 (279)	23.6 (333)
1.51–3.12	0.0	3.6 (408)	7.0 (542)	11.4 (809)
	0.5–1.0	8.1 (125)	11.6 (165)	15.9 (295)
	1.5–5.0	14.8 (44)	24.0 (62)	26.5 (120)

*TOPS Club, Inc., Take Off Pounds Sensibly Club, Inc.
**Prevalence of NIDDM in percent and, in parentheses, subjects per cell.



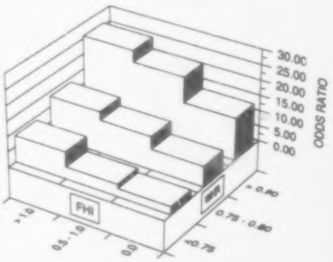


FIGURE 1a—Odds ratios for women between 25 and 50 percent above ideal weight with respect to women less than 25 percent above ideal weight at various levels of family history index (FHI), and waist to hip ratio (WHR) (1969 National Survey of Members of the TOPS Club, Inc)

FIGURE 1b—Odds ratios for women more than 50 percent above ideal weight with respect to women less than 25 percent above ideal weight at various levels of family history index (FHI), and waist to hip ratio (WHR) (1969 National Survey of Members of the TOPS Club, Inc)

the extent provided by calculation of odds ratios of interaction,[16] such an approach is impractical with this number of independent variables.

Results

Table 1 shows that the diabetic women were generally more obese, older, and had a higher FHI than the non-diabetic women. The Pearson correlation between relative weight and WHR was 0.26; WHR has a correlation with age of 0.20, but is only weakly correlated with family history of diabetes ($r = 0.04$).

Table 2 shows the steady increase in diabetes prevalence associated with increases in relative weight and family history. Prevalence rises from 1.4 percent for women in the lowest categories of WHR, relative weight, and family history to 26.5 percent for women in the highest risk group. Odds ratios increase consistently with WHR and family history when the two higher weight categories are compared to the lowest weight group (Figure 1).

Logistic regression results were standardized by calculating the odds ratios associated with an increase of one standard deviation in the independent variables (Table 3). The standardized odds ratio for diabetes associated with the family history index, WHR, relative weight, and age are 1.47, 1.46, 1.54, and 1.14, respectively. The only significant interaction was a standardized odds ratio of 0.94 for the combination of family history and relative weight. Smoking history was not significantly related to diabetes prevalence.

Discussion

The prevalence of diabetes increases with WHR at all levels of family history

TABLE 3—Multiple Logistic Regression Results for Diabetes Prevalence as a Function of Relative Weight, Family History Index (FHI) and Age and Their Interaction among Women from the 1969 National Survey of Members of TOPS Club, Inc*

Variable	Odds Ratio for One Standard Deviation	95% Confidence Interval
Family History Index	1.47	1.38–1.56
Waist:Hip Ratio	1.46	1.42–1.50
Relative Weight	1.54	1.49–1.59
Age	1.14	1.11–1.18
Rel Wt* FHI†	0.94	0.91–.96

*TOPS Club, Inc., Take Off Pounds Sensibly Club, Inc.
†No other interactions were significant at the level, $p < .05$.

and similarly increases with increasing family history for the disease at each level of obesity, and both factors show independent associations with diabetes. These conclusions are supported by the logistic regression results.

An odds ratio less than 1 for the interaction of relative weight and FHI indicates that this interaction is less than multiplicative. No other interactions were significant. This implies that the associations between family history, WHR, relative weight, and age are multiplicative in nature.

Gestational diabetes and undiagnosed diabetes as well as other possible sources of bias such as those associated with variations in awareness of family history, the potential for weight change due to TOPS membership, and the possibility of weight gain or loss following the onset of diabetes apparently do not affect these findings.[16] If present, these biases would tend to reduce the strength of the observed associations. ☐

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Erratum

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All of the column headings in Table 4 (which appeared on page 214) got shifted to the left one column. A corrected version of Table 4 is printed below:

TABLE 4—Percentage of Patients Discharged AMA, by Hospital Type and Major Diagnostic Category

		Teaching Hospitals								
		Suburban		City		Community				
	Major Diagnostic Category	Major	Minor	Major	Minor	Suburban	City	Medical School	Other Specialty	Average
1	Nervous system	0.5	1.1	1.3	1.6	1.0	2.1	1.1	0.0	1.1
2	Eye	0.0	0.0	0.7	0.6	0.1	0.7	0.2	0.0	0.2
3	Ear, nose, and throat	0.0	0.5	0.3	0.4	0.1	0.8	0.4	0.0	0.2
4	Respiratory system	0.6	1.0	1.1	1.3	0.5	2.6	1.2	0.1	1.0
5	Circulatory system	0.6	0.7	1.2	1.0	0.7	2.0	0.8	0.1	1.0
6	Digestive system	0.4	0.6	0.9	0.9	0.5	1.7	0.7	0.0	0.7
7	Hepatobiliary system and pancreas	0.7	0.7	1.8	0.7	1.0	3.7	1.3	0.0	1.4
8	Musculoskeletal system	0.2	0.2	0.6	0.7	0.3	1.1	0.4	0.1	0.5
9	Skin, subcutaneous tissue	0.5	1.1	2.2	1.5	0.5	3.5	1.5	0.0	1.4
10	Endocrine, nutritional and met.	0.5	0.8	1.5	1.1	0.8	2.5	0.8	0.1	1.1
11	Kidney and urinary tract	0.3	0.6	0.8	0.8	0.6	1.2	1.3	0.0	0.8
12	Male reproductive system	0.3	0.1	0.4	0.7	0.3	0.7	0.5	0.4	0.4
13	Female reproductive system	0.0	0.2	0.6	0.3	0.1	0.9	0.4	0.0	0.4
14	Pregnancy, childbirth	0.1	0.0	0.4	0.1	0.2	0.3	0.5	0.0	0.3
15	Newborns and neonates	0.1	0.0	0.0	0.1	0.0	0.0	0.1	0.1	0.1
16	Blood, blood forming and imun.	0.3	2.2	1.6	2.2	1.3	5.2	2.2	0.0	1.7
17	Myeloproliferative, malignancy	0.1	0.0	0.4	0.9	0.1	0.5	0.3	0.1	0.2
18	Infections and parasitic	0.5	0.0	1.5	1.7	0.7	2.8	1.3	0.3	1.1
19	Mental	4.3	4.2	5.0	3.7	7.9	5.3	7.9	1.6	6.4
20	Alcohol/drug use	7.1	14.4	13.4	7.4	12.3	18.7	7.1	0.0	14.6
21	Injuries, poisonings and toxic	2.7	4.9	4.1	6.0	2.6	7.9	1.6	0.2	3.5
22	Burns	0.4	3.6	1.7	1.0	1.0	2.9	3.3	0.0	1.2
23	Factors influencing services	0.9	1.0	0.6	1.4	0.6	3.6	0.3	0.4	0.7
24	Other	0.0	0.0	2.6	0.0	2.0	2.5	0.5	0.0	0.8
	Average	0.55	0.99	1.37	1.00	0.94	3.49	1.0	0.07	1.20

Effects of an Intervention Program on AIDS-Related Drug and Needle Behavior among Intravenous Drug Users

Richard C. Stephens, PhD, Thomas E. Feucht, PhD, and Shadi W. Roman, PhD

ABSTRACT

Background: Behaviors which entail high risk for human immunodeficiency virus (HIV) infection among intravenous drug users can be significantly reduced through educational intervention.

Methods: The educational intervention was conducted by a health educator in a one-on-one format and provided information about acquired immunodeficiency syndrome (AIDS) and strategies for reducing the risk of infection. Risk was assessed prior to the intervention and was compared to a follow-up assessment obtained approximately three months later (n = 322).

Results: The percentage who reported using drugs intravenously decreased from 92.2 to 70.5, and the percentage who reported sharing syringes dropped from 67.4 to 24.3. Reductions of these and other high-risk behaviors were detected across various demographic subgroups, and analyses show that the impact of the intervention endured for several months.

Conclusions: The findings suggest that intravenous drug users conform to a model of rational choice when confronted with accurate information about the devastating consequences of HIV infection. (*Am J Public Health* 1991;81:568–571)

Introduction

Intravenous drug users (IVDUs) constitute the second largest group of persons with acquired immunodeficiency syndrome (AIDS) in the United States and serve as a major conduit of the human immunodeficiency virus (HIV) to other groups, including non-drug-using sex partners and their children.[1,2] Although interviews with successive waves of IVDUs entering drug treatment have suggested some stabilization of HIV seroprevalence as well as some reduction in high-risk behaviors such as sharing needles and the frequency of injection,* these patterns may simply reflect changes in the composition of successive cohorts of IVDUs. Based on repeated interviews with a sample of IVDUs, panel studies have demonstrated some reduction in the frequency of injection and the frequency of sharing needles,[3-5] but the extent to which the reduction of high-risk behavior is directly responsible for the apparent stabilization of seroprevalence rates remains uncertain.[6]

In the past few years, innovative risk-reduction interventions for IVDUs have been developed. Quasi-experimental studies used to evaluate the impact of these interventions have shown risk reduction among IVDUs including a decrease in the frequency of injecting drugs** and an increase in the use of bleach to disinfect shared needles.[7]

Quasi-experimental designs are not without drawbacks. Precautions must be taken to guard against small sample size and low follow-up rate. Since the design may not allow comparisons to any type of control group, other techniques must be used to resolve questions of internal validity. Furthermore, while interviews obtained at as few as two points in time provide the researcher with no direct means of confirming the stability of the change over time, certain data analytic techniques can be used to explore the issue of stability. Each of these drawbacks is addressed in this report in which findings are presented from a quasi-experimental evaluation of an intervention program designed to reduce the chance of HIV infection among IVDUs.

Methods

Subjects

IVDUs recruited between February 1988 and August 1989 for an AIDS education and intervention research program were paid a nominal fee to participate in an initial assessment interview, an education

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*Watters, JK: Preventing human immunodeficiency virus contagion among intravenous drug users: The impact of street based education on risk behavior. Presented at the III International Conference on AIDS, Washington, DC, June 2, 1987.

**Sufian M, Friedman SR, Goldsmith D, Des-Jarlais DC, Delana D: Risk reduction after outreach intervention amongst intravenous drug users. Presented at the V International Conference on AIDS, Montreal, Canada, June 1989.

From the Department of Sociology, Cleveland State University (Stephens, Feucht) and the Cuyahoga County Drug Abuse Services Board, Cleveland (Roman). Address reprint requests to Richard C. Stephens, Department of Sociology, Cleveland State University, Rhodes Tower, Rm 1724, Euclid Avenue at East 24th Street, Cleveland, OH 44115. This paper, submitted to the Journal April 23, 1990, was revised and accepted for publication September 19, 1990.

TABLE 1—Needle and Drug-Related Behavior, Pretest/Posttest Comparisons Two Months Prior to Interview†

	n	Pretest	Posttest	Difference (95% CI)
Needle Behavior				
Scale means				
Using others' works	317	2.28	1.61	43.1 (36.0, 50.2)
Cleaning with bleach	315	1.86	1.87	—
% Sharing works	304	67.4	24.3	43.1 (36.0, 50.2)
% Using others' works	317	92.2	67.1	25.1 (19.4, 30.8)
% Sharing cooker	305	70.3	20.3	50.0 (44.5, 55.5)
% Using drugs IV	322	92.2	70.5	21.7 (16.1, 27.3)
% Cleaning with bleach	278	33.5	62.2	28.7 (21.0, 36.4)
Drug Behavior				
Means				
Number of drugs used	322	4.53	3.30	1.23 (1.13, 1.33)
Number of drugs used IV	322	2.24	1.43	0.81 (0.70, 0.92)
Heroin IV scale	278	4.24	3.32	0.92 (0.77, 1.07)
Cocaine IV scale	311	4.43	3.67	0.76 (0.63, 0.89)
Heroin use scale	322	3.49	2.15	1.34 (1.16, 1.52)
Cocaine use scale	322	4.47	3.27	1.20 (1.04, 1.36)
Speedball use scale	322	1.91	1.06	0.84 (0.70, 0.98)
% using heroin	322	65.5	45.7	25.9 (18.4, 33.4)
% using heroin IV	278	74.5	50.3	24.2 (16.4, 32.0)
% using cocaine	322	83.9	67.2	16.7 (10.8, 22.6)
% using cocaine IV	311	80.4	60.5	19.9 (13.8, 26.0)
% using speedball	322	40.7	27.3	13.4 (6.2, 20.6)

†The sample n varies from 322 due to missing values.

and intervention session, and a follow-up assessment begun at three months after the intervention. Pretest and posttest data were obtained on 402 subjects (60.7 percent of IVDUs from whom pretest assessments were obtained). Approximately 90 percent of the 402 subjects were street addicts not currently in treatment, and the remainder were new admissions to methadone treatment centers. The sample was predominantly male, predominantly Black, and had a median age of 36 years. Eighty subjects are not included in the analyses presented here because they were not considered to be at risk of HIV infection during the time between intervention and posttest (typically because of extended hospitalization, enrollment at a residential drug treatment facility, or incarceration) or because they participated in an alternative intervention. Most subjects were reinterviewed between three and five months after the intervention. However, because some subjects were hard to locate, the amount of time between intervention and posttest interview ranged from 36 to 448 days.

Risk Factor Measures

Pretest and posttest interviews employed multiple measures of two risk factors for HIV infection pertinent to IVDUs. Measures of risk related to needle behavior included sharing works, sharing a cooker, using drugs intravenously, using others' works, cleaning used works with bleach (each recorded as a dichotomy indicating whether the respondent had ever engaged in the behavior), and two scales measuring the frequency of using others' works and cleaning used works with bleach (where 1 = never, 2 = rarely, 3 = 25 percent of the time, 4 = 50 percent of the time, 5 = 75 percent of the time, and 6 = always). Measures of risk associated with more general issues of drug use included: number of drugs used, number of drugs used intravenously, scale scores of heroin use, cocaine use, and speedball use (where the scale is 0 = never, 1 = one to three times a year, 2 = monthly, 3 = one to three times a month, 4 = weekly, 5 = two to three times a week, 6 = four to six times a week, 7 = daily, 8 = two to three times a day, and 9 = four or more times daily), frequency of intravenous use of heroin, intravenous use of cocaine, and intravenous use of speedball (each recorded as the percent of the time that the drug is taken *intravenously* using the percentage scale described above). In addition, each of the variables pertaining to frequency of use or level of intravenous use was recorded as a dichotomy (never/at least once). All questions in the interview regarding these risk behaviors made reference to the two months immediately prior to assessment.

Intervention

The purpose of the quasi-experimental design (pretest-treatment-posttest) was to evaluate the impact of the educational intervention program. This intervention, presented in a one-on-one format by a professionally trained health educator, typically lasted from 45 minutes to an hour and was composed of four modules. The first module presented basic information about the HIV, common modes of transmission, and the progression of the disease. Because the seropositivity rate among IVDUs in Cleveland is quite low (estimates range from 3 to 6 percent based on several studies of selected groups of intravenous drug users conducted in Cleveland by the Cuyahoga County Drug Abuse Services Board) and consequently few of our clients had any immediate acquaintance with AIDS, this module centered on a segment from the film "AIDS: A Bad Way to Die,"*** produced by prisoners who were dying of AIDS. The second and third modules focused on reducing risks associated with sexual behavior and drug behavior. Condom use was demonstrated, and the relative risks of various sexual practices were discussed. Subjects were advised to modify their drug behavior by halting their use of drugs, not using drugs intravenously, not sharing works, or (as a defense of last resort against HIV infection) cleaning works when sharing with others and were shown how to clean a needle and syringe with bleach. Each subject was provided with a kit of materials including a bottle of bleach, condoms, and various brochures about AIDS. In the final module, subjects were told about the HIV antibody test and what both positive and negative results meant and were encouraged to take advantage of the anonymous blood testing provided as part of the project.

Evaluation of the intervention focused on three research questions:

- Do comparisons of pretest and posttest scores of risk behaviors provide evidence that the intervention was successful in reducing the level of risk for HIV infection among the respondents?
- In the absence of a true control group, can the impact of the intervention be assessed using a non-equivalent comparison group?
- Did the intervention have a sustained effect?

***A production of the Taconic Video Center, Taconic Correctional Facility, Taconic, NY, 1986.

TABLE 2—Assessment of Duration of Impact of the Intervention: Mean Values on Risk Behaviors for Different Pretest-Posttest Intervals

Length of Pretest-Posttest Interim	Group I ≤ 111 days n: 101	Group II 112–154 days n = 101	Group III ≥ 155 days n = 101	F-ratio ($p < 0.05$)
Needle Behavior				
Scale Means				
Using other's works	1.65	1.75	1.44	—
Cleaning w/bleach	1.70	2.12	1.80	—
% Using other's works	58.3	77.6	65.4	12.051
% Sharing works	19.6	29.0	20.6	—
% Sharing cooker	26.6	45.3	28.1	—
% Using drugs-IV	65.7	80.4	65.4	7.519
% Cleaning w/bleach	96.4	100.0	94.2	—
Drug Behavior				
Scale Means				
# Drugs used	3.11	3.57	3.22	—
# Drugs used-IV	1.21	1.69	1.39	3.987
Heroin use scale	1.66	2.36	2.45	—
Heroin IV scale	2.56	3.33	3.37	4.090
Cocaine use scale	3.41	3.60	2.82	—
Cocaine IV scale	3.59	3.99	3.19	3.229
Speedball use scale	.34	.50	.47	—
% Using heroin	35.2	51.4	50.5	7.196
% Using heroin-IV	32.4	53.3	50.5	11.139
% Using cocaine	66.7	75.7	60.7	—
% Using cocaine-IV	55.6	70.1	50.5	9.16
% Using speedball	21.3	31.8	29.0	—

Results

Six of the seven needle behavior risk behaviors decreased in frequency at the posttest (Table 1). A major reduction was obtained in the percentage who reported sharing their works, and a considerable increase was obtained in the percentage who reported cleaning works with bleach. All 12 measures of risk associated with general drug behavior recorded significant risk reduction. No significant differences in risk reduction were found between IVDUs newly enrolled in treatment and those not in treatment, or between gender or ethnic groups. Additional analyses revealed no systematic variation among groups of IVDUs in terms of knowledge about AIDS at the time of pretest.

Comparison of Experimental and Control Groups

Since a true control group was ethically unacceptable, a non-equivalent control group was constructed composed of individuals who were interviewed for the pretest at approximately the same time as those interviewed for the posttest. For 19 of the 21 measures of risk associated with needle behavior, the "experimental" (posttest) group reported significantly lower levels of risk than the "control" (contemporaneous pretest) group. Similar contrasts were found in reduction of risk associated with general drug behaviors. (Details of these comparisons are available from the authors.)

Duration of the Impact

The important question of whether the intervention had a lasting impact on reducing high-risk behavior is difficult to answer directly since interviews were conducted at only two points in time for each subject. Fortuitously, indirect evidence can be obtained using the variation in the length of time between the intervention and the posttest interview. Subjects were divided into three equal-sized groups: those interviewed within 111 days of the intervention, those interviewed between 112 and 154 days following the intervention, and those interviewed more than 154 days after intervention. Comparisons of these groups in terms of the magnitude of risk reduction are presented in Table 2. While the results do not clearly establish the stability of the intervention's impact, neither do they suggest a significant decay over time. As in the previous analyses, a greater impact was observed for needle-related behaviors than for general drug use behaviors.

Discussion

The data demonstrate that significant changes occurred between pretest and posttest in needle and drug-related behaviors of intravenous drug users, and comparisons of experimental groups with contemporaneous non-equivalent control groups suggest that these changes are at least partly attributable to the intervention. In addition, the data indicate that the impact of the intervention endured for up to one year. Finally, while reductions were detected in a broad range of behaviors associated with a high risk for HIV infection, analyses suggest that the intervention had a greater impact upon specific needle-related behaviors (such as sharing works or cleaning works) than upon more general drug-related behaviors (such as frequency of heroin or cocaine use.) All of these effects were achieved across gender and ethnic groups. These data replicate many of the findings of other earlier panel studies.[3,4,7]

Several specific limitations of the study should be noted:

- Measures of risk are based exclusively on self-report. A number of studies indicate that addicts in research projects are usually honest respondents,[8–10] and interviewers were trained to elicit the subjects' confidence. The hour-long interview was fairly complicated making misrepresentations difficult to maintain. Moreover, the variability in findings in this report (and the generally less favorable impact of the intervention upon sexual behavior to be reported elsewhere) convinces us that our data are generally reliable.
- The use of convenience sampling complicated by self-selection bias and attrition weakens generalizations to a population. Although it does not completely dispel concerns about sample bias, a comparison of dropouts and those subjects who completed the study revealed almost no significant differences in terms of pretest risk behaviors, gender, ethnicity, or age. (These data are available from the authors.) This suggests that the final study sample adequately reflects the original sample of all subjects at pre-test.
- The lack of a true control group lessens the extent to which issues of internal validity can be adequately addressed.[11]

Despite these shortcomings, these findings suggest that many addicts act in a rational fashion when faced with important choices regarding behaviors which are directly AIDS-related. It appears that the educational intervention was successful because it sensitized the subjects to the realities of AIDS, established AIDS as a real threat to them, and then immediately armed them with knowledge and materials (bleach, condoms, and assorted literature) to assist them in their own risk-reduction efforts. Levels of drug use seem somewhat resistant to intervention in this sample. However, when faced with the prospect of disease and death from AIDS, many addicts seemed willing to modify their behavior—particularly the ways in which needles and related paraphernalia are used—in order to minimize the chance of infection. ☐

Acknowledgments

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Harbor-UCLA Selected for Psychiatry Research Center

The National Institute of Mental Health (NIMH) recently approved the establishment of the nation's first mental health research center, the Research Center of the Psychobiology of Ethnicity, at the Harbor-UCLA Medical Center in Torrance, California. The $1.5 million grant will be administered by Harbor-UCLA's Research and Education Institute.

Alan I. Leshner, PhD, Acting Director of NIMH, said, "Although mental illness can strike anyone regardless of race or cultural origin, there appear to be some very important differences in the ways people of different ethnic backgrounds respond to medications in the treatment of depression, schizophrenia and other serious mental disorders. The research undertaken by the Harbor-UCLA Hospital Center will greatly enrich our understanding of the psychobiological bases of mental disorders and their treatment, and help ensure that every patient of every ethnic background receives appropriate treatment."

Center director Keh-Ming Lin, associate professor of psychiatry at Harbor-UCLA, described the focus of the center's programs as multidimensional and designed to serve the multiethnic populations of the community. The center will study the role that race and ethnicity play in the differing responses to treatments of serious mental illnesses. Drugs used in the management of mental depression, mania, and schizophrenia have been shown to have differing impacts, and to reach different levels in the blood and the brain, with ethnicity being a significant variable.

In addition to the study of biological issues and their impact on mental illnesses in people of differing ethnicity, the center will also study how cultural factors influence the evaluation and care of patients, and how they impact on the willingness of patients and families to make use of treatments and support programs found useful in mental health care.

Lin explained that "a very serious problem exists among minority patients and their families because they rarely make use of the self-help and patient/family support groups like the California Alliance for the Mentally Ill which provide so much assistance to those who do participate."

Lin and his colleagues are currently involved in a research project with Hispanic patients and families to identify ways to make treatment programs and self-help group activities more culturally acceptable and to increase access for these individuals to such programs. Other studies underway focus on the effect of mental illnesses on Asian, Black, and Native American patients and their families.

For additional information, contact the Office of Scientific Information, National Institute of Mental Health, 5600 Fishers Lane, Room 15-105, Rockville, MD 20857; (301) 443-4536.

Adult Survivors of Childhood Sexual Abuse and Subsequent Risk of HIV Infection

Sally Zierler, DrPH, Lisa Feingold, MSPH, Deborah Laufer, MSPH, Priscilla Velentgas, Ira Kantrowitz-Gordon, ScB, and Kenneth Mayer, MD

ABSTRACT

Background: Epidemiologic description of long-term adverse health effects of childhood sexual abuse is lacking, despite estimates that perhaps 30 percent of adults have experienced sexual assault in childhood.

Methods: In an adult cohort enrolled to investigate causes of transmission of human immunodeficiency virus, we identified current behaviors affecting risk of infection that were associated with a history of early sexual abuse. One hundred and eighty-six individuals provided information on the occurrence of abuse and subsequent sexual and drug using activities.

Results: Approximately half of the women and one-fifth of the men reported a history of rape during childhood or adulthood. Twenty-eight percent of the women and 15 percent of the men recalled that they had been sexually assaulted during childhood. People who reported childhood rape compared with people who did not were four times more likely to be working as prostitutes (90 percent confidence interval = 2.0, 8.0). Women were nearly three times more likely to become pregnant before the age of 18 (90% CI = 1.6, 4.1). Men who reported a history of sexual abuse had a twofold increase in prevalence of HIV infection relative to unabused men (90% CI = 1.0, 3.9).

Conclusions: The disturbing prevalence of early sexual abuse and its possible health-related consequences call for prompt and routine investigation of sexual abuse histories. Identification of sexual victimization may be an important component for management of risk factors for human immunodeficiency virus. (*Am J Public Health* 1991;81:572–575)

Introduction

The social sciences have become increasingly concerned with the prevalence and psychosocial consequences of sexual abuse in childhood.[1-10] Prevalence studies in communities of free-living adults estimate that 6 to 62 percent of women and 3 to 31 percent of men were sexually abused before the age of 18.[11] The range in these estimates is a function of the specificity of the definition of sexual abuse and the age definition for childhood. When the definition is limited to children under 14 years of age who reported physical contact of a sexual nature, estimates range from 28 to 36 percent.[3] In 1985, the American Humane Association reported 123,000 cases of childhood sexual abuse during that year alone, and yet this figure is likely to be a gross underestimate.[11]

Childhood sexual abuse is a medical problem that requires careful diagnosis and follow up for trauma,[12] pregnancy,[13] disease transmission, including human immunodeficiency virus (HIV) infection,[13-16] and acute psychological morbidity.[11-13,17,18] Two behavioral outcomes that are associated with sexual abuse in childhood are chemical substance abuse and prostitution.[10,11] These risk-taking behaviors have health-related consequences that theoretically could include any diseases resulting from intravenous drug use or unprotected sexual activity with multiple partners. Infection from HIV would be associated with childhood sexual abuse if such abuse predisposes children to drug using or sexual behaviors that involve contact with bodily fluids infected with this virus.

An explicit goal of our study of determinants of heterosexual transmission of HIV infection in the Rhode Island area was to learn if a history of sexual abuse is a risk factor for subsequent HIV infection in adulthood. We hypothesize that adults who were sexually abused as children are more likely to participate in activities that increase their risk of exposure to HIV.

Methods

Study Population

All women and men at least 18 years old, residing or attending school in southeastern New England, currently or recently sexually active with someone of the opposite sex, and who signed an informed consent statement to participate in the research protocol were eligible to be in the study. Potential study volunteers contacted the study clinic as a result of the recommendation from a health care provider (provider-referred) or because of interest stimulated by a large scale publicity campaign (self-referred).

From the population of 567 women and men who enrolled in the testing and counseling program between March 1988 and October 1989, 393 people were invited to participate in the longitudinal study because they had a nonzero risk of acquiring or transmitting HIV infection (either because they had multiple sexual partners, were themselves HIV-infected, or had

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From the Department of Community Health, Division of Biology and Medicine, Brown University, all authors except Laufer who is with the Infectious Disease Division at Memorial Hospital of Rhode Island where Dr. Mayer is also affiliated. Address reprint requests to Sally Zierler, DrPH, Department of Community Health, Brown University, Box G-A405, Providence, RI 02912. This paper, submitted to the Journal May 2, 1990, was revised and accepted for publication October 10, 1990.

TABLE 1—Demographic Characteristics according to Reported History of Childhood Sexual Abuse

	Child Sexual Abuse* (n = 41)	No Sexual Abuse (n = 123)	Total (n = 164)
	%	%	%
Gender			
Female	71	44	51
Male	29	56	49
Age (years)			
18–24	10	16	15
25–34	58	51	53
35–44	27	26	26
45+	5	7	6
Race			
White	76	80	79
Other	24	20	21
Referral			
Provider	54	47	49
Self	46	53	51

*Reported history of sexual abuse as a child or teenager.

TABLE 2—Prevalence of Reported Sexual Abuse according to Time of First Occurrence

	Women (n = 101)		Men (n = 85)		Total (n = 186)	
	n	%	n	%	n	%
Never sexually abused	54	53.5	69	81.2	123	66.1
Childhood	20	19.8	8	9.4	28	15.1
Teenager	9	8.9	4	4.7	13	7.0
Adult*	18	17.8	4	4.7	22	11.8

*22 individuals reporting first abuse as adults were deleted from all subsequent analyses.

partners who were infected or at increased risk for HIV infection). Selection of the cohort and the decision to participate were independent of a history of childhood sexual abuse, since a history of sexual abuse was not known by the study personnel at the time of enrollment, nor had the potential participant seen or completed the questionnaire which investigated that history. There were 206 people who were eligible for the longitudinal study but refused to participate. The discontinued study population relative to the continuing study population was more likely to be male and less likely to be HIV-infected. The people who discontinued compared with those who agreed to remain were similar with respect to their age distribution, number of different sexual partners, race and ever use of intravenous drugs. Among the 187 individuals who agreed to continue in the study by signing a second informed consent statement, 186 had complete information on history of sexual assault. This population provided the information for the current study.

Exposure Information: A certified HIV counselor or nurse-clinician administered a lengthy behavioral questionnaire to study participants at the time they agreed to be in the longitudinal study. Information on the exposure of interest, a reported history of sexual abuse, was based on the question, "Have you ever been raped or forced to have sex?" People who acknowledged a history of sexual abuse further remarked on the frequency and timing of this occurrence (during childhood, adolescence, adulthood or during any combination of these time periods).

For the purpose of this analysis, individuals were defined as exposed to childhood sexual abuse if they reported a history of rape or forced sex at least once during childhood or as a teenager. Individuals were defined as not exposed if they reported no history of rape or forced sex, including during adulthood. Individuals reporting a history of first rape as an adult (n = 22) were excluded from the analyses due to our inability to determine a temporal sequence between exposure to sexual abuse and the outcomes of interest.

Outcome Information: Outcomes of interest included behaviors that may increase risk of acquiring HIV infection, as well as biologic parameters such as a history of sexually transmitted disease or current HIV infection. Information for the behavioral outcomes was obtained from the questionnaire administered at entry into the longitudinal study. Individuals provided information on their lifetime use of intravenous and non-intravenous illicit drugs, alcoholic beverages, contraceptives, and prior or current work as a prostitute. Detailed history of number of sexual partners, number of casual sexual partners, number of years of sexual activity, sexual practices, education, and pregnancy experience provided additional information on relevant outcomes.

Sexually transmitted disease history was obtained by self-report. HIV antibody status was determined by serologic testing by standard enzyme-linked immunoassay and Western Blot procedures using fresh blood samples that were drawn at the time of entry into the study.

All outcome measurements were created as dichotomous variables defined as the presence or absence of the outcome of interest. Examination of non-intravenous drug use and alcohol use was limited to those reporting no history of intravenous drug use because of the strong association of intravenous drug use with nonparenteral use of other drugs.

Data Analysis: Prevalence ratios provided a quantitative measure of risk between the outcome of interest and child sexual abuse. Prevalence ratios were computed as the proportion of people who had a particular outcome within the population who reported a history of child sexual abuse relative to the proportion who had this outcome within the population who reported no history of sexual abuse at any time.

Multiple logistic regression was used to examine the possible confounding effects of gender, race, and study referral source. Evaluation of confounding indicated that race and referral pattern did not materially change the coefficient for the child abuse term (for all outcomes, the coefficient for child abuse changed less than 5 percent) and therefore these potential confounders were deleted from the analysis. Table 1 lists the demographic characteristics of the cohort that was included in the analysis.

Because gender was the only measured confounder, and also because of in-

TABLE 3—Behaviors Associated with a History of Childhood Sexual Abuse

	Childhood Sexual Abuse							
	Women			Men			Overall	
Behavior	Yes (n = 29)	No (n = 54)	Prev. Ratio	Yes (n = 12)	No (n = 69)	Prev. Ratio	Prev. Ratio$_{MH}$*	90% Confidence Interval
Prostitution	6	4	2.8	4	3	7.6	4.0	(2.0, 8.0)
IVDU**	8	13	1.2	6	32	1.2	1.2	(.78, 1.8)
Average yearly sexual partners (>2 vs ≤2)	15	16	1.7	10	24	2.4	2.0	(1.4, 2.8)
Never/rarely use condom in last year†	21	36	1.0	5	41	0.8	1.0	(.76, 1.2)
Sex with strangers††	23	28	1.5	10	52	1.2	1.4	(1.1, 1.6)
Education (<HS vs ≥HS)	9	5	3.4	2	15	0.8	1.9	(1.0, 3.5)
Pregnancy before 18 years old	18	13	2.6	—	—	—	2.6	(1.6, 4.1)

*Prev ratio$_{MH}$ are Mantel-Haenszel estimates of prevalence ratios pooled over gender strata.

**One woman and one man with history of child sexual abuse were deleted from analysis because temporal sequence of aga of sexual abuse and age of IV drug use not specified

†Four sexually inactive individuals deleted from analysis.

††Eight individuals deleted due to missing data.

TABLE 4—Use of Alcohol and other Drugs and Childhood Sexual Abuse among Adults with no History of Intravenous Drug Use*

	Childhood Sexual Abuse							
	Women			Men			Overall	
Drug use†	Yes (n = 20)	No (n = 41)	Prev. Ratio	Yes (n = 5)	No (n = 37)	Prev. Ratio	Prev. Ratio$_{MH}$**	90% Confidence Interval
Cocaine	8	20	0.8	3	20	1.1	0.9	(0.59–1.3)
Tranquilizer	9	11	1.7	2	9	1.6	1.7	(0.98–2.8)
Marijuana	19	33	1.2	4	32	0.9	1.1	(0.95–1.3)
Hallucinogens	6	14	0.9	1	10	0.7	0.8	(0.46–1.6)
Heavy alcohol	6	6	2.0	3	14	1.6	1.8	(1.0–3.3)

*22 women and 39 men intravenous drug users deleted from analysis.

**Prev ratio$_{MH}$ are Mantel-Haenszel estimates pooled over gender strata.

†Drug use was defined as ever use of these drugs (no use was defined as never used); Heavy alcohol use was defined as usual consumption of at least one bottle of wine, five cans of beer, or one pint of hard liquor when drinking (no heavy use was typically drinking less than this amount.)

terest in describing gender-specific associations, all analyses were stratified by gender and summarized as Mantel-Haenszel summary prevalence ratios with 90 percent test-based confidence intervals.[19]

Results

Approximately one out of every two women and one out of every five men reported a history of rape or forced sex (Table 2). The proportion of women and men who reported sexual abuse during childhood or adolescence was 28 and 15 percent, respectively.

The associations for behavioral characteristics and exposure to early sexual abuse appear in Table 3. Survivors of child sexual abuse were four times more likely to report having worked as a prostitute during their lifetime than were those who reported no history of abuse. This higher prevalence of prostitution among survivors was seen among both women and men. Male survivors of abuse were nearly eight times more likely to report a history of prostitution. All of the men reporting prostitution also reported bisexual experiences. Excluding male prostitutes, bisexuality among men was not associated with childhood sexual abuse (prevalence ratio = 1.1). Additional analyses, simultaneously controlling for gender and intravenous drug use, a behavior that may lead to prostitution as a source of income for drugs, resulted in only a slight lowering of the point estimate (prevalence ratio = 3.6, 90 percent CI = 1.8, 7.3). Use of intravenous drugs was only weakly associated with a history of early sexual abuse.

Penetrative sex without a condom with a partner suspected of having HIV infection was similar among survivors of child sexual abuse and among those not abused (prevalence ratio = 1.1). No difference between the two groups existed with respect to the lack of condom use in the year prior to enrolling in the study. Survivors of sexual abuse, however, reported a 40 percent excess frequency of a history of sex with someone they did not know, and were two times more likely to have multiple sexual partners on an average yearly basis relative to people reporting no sexual abuse.

Women who survived sexual abuse had greater than a three-fold increase in risk of not completing high school than women who reported no sexual abuse. This increase was explained in part by a higher occurrence of teenage pregnancy among the women who did not finish high school. Adjustment for pregnancy prior to the age of 18 lowered the point estimate for the relation between education level and sexual abuse history from 3.4 to 1.9 (90 percent CI = 0.8, 4.3). Teenage pregnancy was more prevalent among female survivors of sexual abuse than among those reporting no history of abuse (prevalence ratio = 2.6, 90 percent CI = 1.6, 4.1).

Table 4 describes drug use associations among people with no history of intravenous drug use. Survivors of sexual abuse had a 70 and 80 percent excess of use of tranquilizers and heavy alcohol use, respectively, relative to people reporting no sexual abuse. Women were two times more likely to be heavy consumers of alcohol if they had experienced sexual abuse in childhood than women who did not report this experience.

Overall, survivors of child sexual abuse were only slightly more likely to report a history of sexually transmitted disease (gender-adjusted prevalence ratio = 1.3, 90 percent CI = 0.93, 1.9) or to have serologic evidence of HIV infection (gender-adjusted prevalence ratio = 1.3, 90 percent CI = 0.82, 2.0). However, men who reported early sexual abuse had a two-fold increase in prevalence of HIV infection relative to unabused men (90 percent CI = 1.0, 3.9). This increased prevalence of HIV infection among men who survived sexual assault was not explained by a history of intravenous drug use.

Discussion

National estimates are that one in four girls and one in six boys are sexually assaulted before the age of 18.[20] Over one-fifth of our cohort acknowledged that they had been raped or forced to have sex during childhood or adolescence. This experience was twice as frequent among women as among men.

We do not have evidence of confounding of sexual abuse effects. Our observations are consistent with other studies that have described sexual abuse to be independent of race.[4,11] Unidentified correlates of childhood sexual abuse that are independent risk factors for behaviors leading to HIV infection could explain the positive results in this study. However, we did not ask about child and adolescent experiences other than those related to sexual and reproductive experiences and age of first use of illicit drugs. Childhood sexual abuse, particularly if the perpetrator is a family member, may be accompanied by other forms of abuse and neglect.[9] Our study did not investigate characteristics of family dysfunction that may be associated with both sexual abuse and adult behaviors leading to HIV infection.

We have evidence that early sexual abuse is associated with behavioral outcomes that may be having devastating effects on the public health, particularly in relation to the HIV epidemic. In our study, sexually abused women and men were more likely to engage in sex work, to change sexual partners frequently, and to engage in sexual activities with casual acquaintances than people who were never sexually abused. Women survivors of sexual assault reported more frequent use of large quantities of alcoholic beverages and both genders used tranquilizers more frequently than individuals who had never been assaulted.

The disturbing prevalence of early sexual abuse and its possible consequences on increasing behaviors that could lead to HIV infection and other poor health outcomes have implications for medical and public health practitioners. The safer sex messages, for example, may be missing the point for people whose lives have been complicated by sexual victimization. Social scientists have urged for early identification of children and adolescents who may have been sexually assaulted because of evidence that the earlier the child can begin recovery from this trauma, the better the prognosis for normal adult functioning[3,18] The high frequency of occurrence of sexual assault is not only an immediate threat to the health of children, but may continue to disrupt development of appropriate health behaviors that prevent disease in adulthood. □

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Preventing Homicide: An Evaluation of the Efficacy of a Detroit Gun Ordinance

Patrick W. O'Carroll, MD, MPH, Colin Loftin, PhD, John B. Waller, Jr., DrPH, David McDowall, PhD, Allen Bukoff, PhD, Richard O. Scott, James A. Mercy, PhD, and Brian Wiersema

ABSTRACT

Background: In November 1986, a Detroit, Michigan city ordinance requiring mandatory jail sentences for illegally carrying a firearm in public was passed to preserve "the public peace, health, safety, and welfare of the people."

Methods: We conducted a set of interrupted time-series analyses to evaluate the impact of the law on the incidence of homicides, hypothesizing that the ordinance, by its nature, would affect only firearm homicides and homicides committed outside (e.g., on the street).

Results: The incidence of homicide in general increased after the law was passed, but the increases in nonfirearm homicides and homicides committed inside (e.g., in a home) were either statistically significant or approached statistical significance (p = .006 and p = .070, respectively), whereas changes in the incidence of firearm homicides and homicides committed outside were not statistically significant (p = .238 and p = .418, respectively). We also determined that the ordinance was essentially unenforced, apparently because of a critical shortage of jail space.

Conclusions: Our findings are consistent with a model in which the ordinance had a dampening effect on firearm homicides occurring in public in Detroit. The apparent preventive effect evident in the time series analyses may have been due to publicity about the ordinance, whereas the small nature of the effect may have been due to the lack of enforcement. (*Am J Public Health* 1991;81:576–581)

Introduction

The homicide rate in Detroit, Michigan, was the highest among the 15 largest United States cities in all but three of the last 15 years.[1,2] In 1986, the crude homicide rate in Detroit was 59 per 100,000, the highest of any large city in the United States. As is the case nationally, most homicide victims in Detroit were killed with firearms. In 1985 in Michigan, firearm homicides accounted for 72.7 percent of the homicide-attributable premature loss of life among Blacks. Two-thirds of Michigan's homicides occurred in the Detroit area.[3]

On November 26, 1986, the Detroit City Council, responding to a wave of youth homicides, enacted an ordinance[4] that imposed a mandatory jail sentence on anyone convicted of unlawfully concealing a pistol or carrying a firearm in the city. Although the ordinance contained some technical amendments to the city ordinances regulating the carrying and selling of firearms, the key element of the ordinance was the mandatory sentence of 30 to 90 days in the Wayne County jail and a fine of $100 to $500, depending on whether or not it was a first conviction under the ordinance. The ordinance became effective on January 10, 1987.

One of the primary goals of this legislation was to improve the health of the citizens of Detroit by decreasing firearm-related injuries and deaths. It was specifically stated in the ordinance that this legislation was "necessary for the preservation of the public peace, health, safety, and welfare of the people of the City of Detroit." This ordinance may therefore be viewed as a public health intervention, and its efficacy evaluated in that context.

Underlying this ordinance is the hypothesis that, by decreasing the number of persons carrying firearms on the streets of Detroit, the incidence of firearm homicides resulting from interpersonal violence occurring "outside" (e.g., on streets, in alleys, or in parking lots) would likewise decrease. Since the ordinance does not address the issues of possessing or keeping firearms in private places (such as the home) or of carrying other weapons (such as knives) in public, the law should have its greatest effect on the incidence of firearm homicides and homicides committed in public places while having little effect on the incidence of nonfirearm homicides or on firearm homicides committed inside private residences and other buildings.

To determine whether the imposition of the mandatory jail sentences in January 1987 was associated with changes in the incidence both of homicides committed with firearms and homicides committed outside, we analyzed homicide incidence data in the City of Detroit for the years

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From the Division of Injury Control, Center for Environmental Health and Injury Control, Centers for Disease Control, US Public Health Service, Atlanta, GA (O'Carroll, Mercy); Institute of Criminal Justice and Criminology, University of Maryland, College Park, MD (Loftin, Wiersema); Center for Prevention and Control of Interpersonal Violence, Department of Community Medicine, Wayne State University School of Medicine, Detroit, MI (Waller, Bukoff, Scott); and School of Criminal Justice, State University of New York, Albany, NY (McDowall).

Address reprint requests to Patrick W. O'Carroll, MD, MPH, Mailstop F36, Centers for Disease Control, 1600 Clifton Road NE, Atlanta, GA 30333. This paper, submitted to the Journal August 6, 1990, was revised and accepted for publication December 20, 1990.

1980 through 1987. For comparison, we also analyzed nonfirearm homicides and homicides committed inside residences and other buildings. We predicted that changes attributable to the 1987 ordinance would be reflected in the incidence of firearm homicides and homicides committed outside, but not in the incidence of nonfirearm homicides or homicides committed inside.

In addition, we evaluated the enforcement of the mandatory sentencing ordinance by the criminal justice systems in Detroit. Although one might assume that the ordinance would increase the average sentence for persons convicted of illegally carrying a firearm in Detroit, experience in Detroit and other locations indicates that mandatory sentences do not necessarily produce the expected changes in average sentences.[5-9] For a variety of reasons, judges find ways to circumvent the mandatory sentences, or discretion for sentencing moves to another point in the criminal justice system such as the prosecutor or the police. An examination of how the law was implemented is therefore essential to the interpretation of the time-series analysis. In addition, the Detroit experience is important for what it may reveal about how a criminal justice system adapts to mandatory sentencing for carrying a firearm illegally.

Methods

Evaluating Effect of the Ordinance

Homicide data for the years 1980 to 1987 were drawn from computerized Official Complaint Report (OCR) records maintained by the Detroit City Police Department. All criminal homicides were included in the analysis, including offenses coded as murder, murder (1st degree), murder (2nd degree), and manslaughter.

OCR codes were used to divide the homicides two ways: 1) those in which the victim was killed with a *firearm* (handgun; shotgun/rifle) or by *some other means* (bodily force; black jack; knife, dirk, razor/other sharp objects; blunt object; miscellaneous other); and 2) those in which the victim was killed *outside* (on street; off street, alley; playground/park; parking lot; in a vehicle; vacant lot) or *inside* (building; housing project). Four cases in which the homicide method was coded as "missing" and 73 cases coded as "unknown" were assigned to the nonfirearm series on the grounds that firearm homicides are easily recognized. The police data did not permit a determination of whether homicides in "buildings" occurred in private buildings (such as in a home) or in public buildings (such as in a bar or convenience store). Including homicides in public buildings in our *outside* series would have been preferable, since weapons used to commit such homicides would presumably have been carried "outside" to these buildings, and thus the perpetrators should theoretically have been affected by the ordinance. Nevertheless, if a substantial proportion of the homicides in buildings occurred in public buildings, this would result in a bias toward the null in our study since it would tend to dampen the hypothesized preventive effect on homicides committed outside relative to those committed inside.

Inferences about the impact of the ordinance on homicides are based on an interrupted time-series research design that contrasts the monthly incidence of certain types of homicide before and after the ordinance took effect. A change in the behavior of a particular time series at the point of intervention would provide evidence that the ordinance had an effect on that time series. To measure the size and statistical significance of time-series changes associated with the implementation of the gun ordinance, we used the intervention analysis developed by Box and Tiao based on the Box and Jenkins autoregressive, integrated, moving average (ARIMA) modeling of time series.[10,11] In general terms, the procedure starts with the development of a *noise model* to account for systematic components—trend or drift, seasonality, and autocorrelation in the time series. Modeling these systematic components to explicitly remove their effects is important because it reduces the risk that these components will be confounded with an intervention effect. This noise model essentially controls for variables which influence the homicide rates consistently over time, allowing analysts to estimate the effects of the intervention. After an appropriate noise model to control the effects of these series characteristics has been specified, an *intervention* model is added to represent the effect of the ordinance. If there is a change in the series at the point of intervention, the intervention model will measure the size of the effect.

Ideally, we would have conducted the analysis on monthly observations for each of the four time series resulting from the logical intersection of the two dimensions of interest: 1) firearm homicides in public places; 2) nonfirearm homicides in public places; 3) firearm homicides in nonpublic places; 4) nonfirearm homicides in nonpublic places. However, there were insufficient cases to permit ARIMA modeling of the time series in this manner. Therefore, we conducted two separate monthly analyses: one for firearm versus nonfirearm homicides (categories 1 + 3 versus 2 + 4) and one for public versus nonpublic homicides (categories 1 + 2 versus 3 + 4).

All estimates were computed using the backforecasting algorithm implemented in BMDP2T software.

Evaluation of Enforcement

To assess the impact of the mandatory jail sentencing ordinance on the processing of gun ordinance violators in the Detroit criminal justice system, we monitored the disposition of cases charged under the ordinance during 1987 and 1988 using data that are routinely collected by the prosecutor and the court. In addition, we conducted non-structured interviews with key people in the Detroit criminal justice system, people whose jobs involved implementing the ordinance or processing defendants charged with violating the ordinance, e.g. members of the Detroit Police Department, the Department of the Wayne County Prosecuting Attorney, the Office of the Detroit Corporate Council (city solicitor), the 36th District Court, and others. These interviews were not intended to constitute a scientific survey; rather, they were intended to supplement the statistical analysis of the disposition of offenders charged under the ordinance by allowing a description of perceptions regarding the enforcement of the ordinance by those most likely to be well informed in this regard.

Results

Findings of the Time-Series Analysis

The two sets of time series data—firearm vs nonfirearm homicides, and inside vs outside homicides—are plotted in Figures 1 and 2. The results of the interrupted time-series analyses of these data are summarized in Table 1. For each of the four series that we analyzed, the table shows the estimated size of the intervention effect (in terms of the change in the number of homicides per month associated with the intervention); the mean number of homicides per month before the intervention; the relative magnitude of the intervention effect (expressed as a percent change in the preintervention mean number of homicides per month); and the

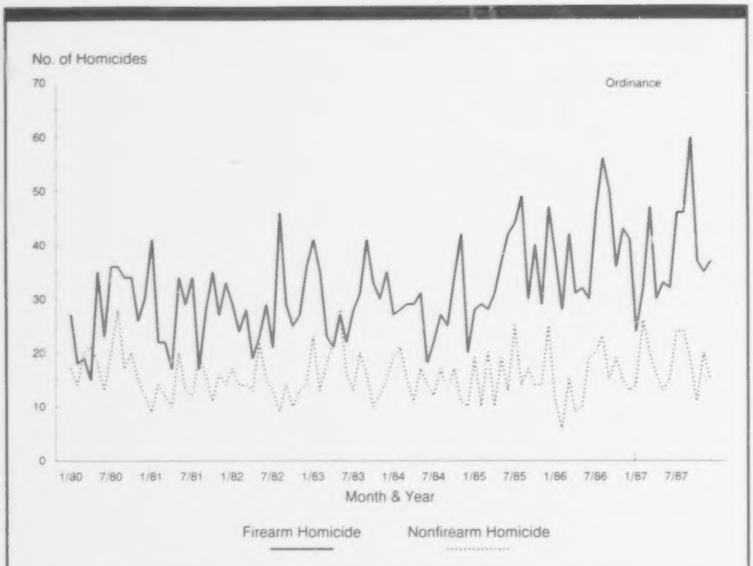

FIGURE 1—Number of Homicides in Detroit by Month and Weapon Used, 1980–87

p-value, representing the probability of a Type-I error associated with a test of the null hypothesis of a zero intervention effect. In the Appendix, we provide more complete information on the time-series analysis.

The clearest feature of the analysis, reflected in the positive estimates of the intervention effect in all four series (Table 1) is that in January 1987, as the ordinance was implemented, Detroit was experiencing a substantial increase in homicides. This is consistent with the public perception in Detroit of increasing homicide incidence. As noted above, this perception is what stimulated the Council to pass the mandatory sentencing ordinance.

Although the ordinance clearly failed to stop the surge, a comparison among the various time series of changes associated with the intervention in the mean number of homicides per month suggests that the ordinance may have had a dampening effect on the increasing incidence of homicide. Homicides committed inside (those unlikely to be affected by the ordinance) increased by 22 percent, a much greater increase than the 10 percent increase in homicides committed outside (those more likely to be affected by the ordinance). The statistical analysis shows that the increase in inside homicides is large relative to its standard error (see Appendix) and is therefore unlikely to be due to a chance fluctuation (p = .006); the increase in outside homicides, however, is small relative to its standard error and is compatible with a variation attributable to chance (p = .418). The data are therefore consistent with a model in which the incidence of homicides committed inside increased, but the incidence of homicides committed outside was unchanged.

A similar but less pronounced pattern exists for the firearm/nonfirearm contrast. The incidence of nonfirearm homicides increased 16 percent after the intervention, a change that is large relative to the standard error of the series and therefore relatively unlikely to be due to a chance fluctuation (p = .07). Although the incidence of firearm homicides increased almost as much (13 percent over the mean of the preintervention series), the change was small relative to its standard error and not statistically significant by conventional standards (p = .24). In other words, the increases in the two types of homicide most clearly targeted by the ordinance were not statistically significant; plausibly the increases observed were random fluctuations around the preintervention mean. In contrast, the increases in the two comparison series that were least likely to be influenced by the ordinance either were clearly statistically significant or approached statistical significance at the p = .05 level. The patterns are what would be expected if the Detroit gun ordinance had a dampening effect on firearm homicides and outside homicides.

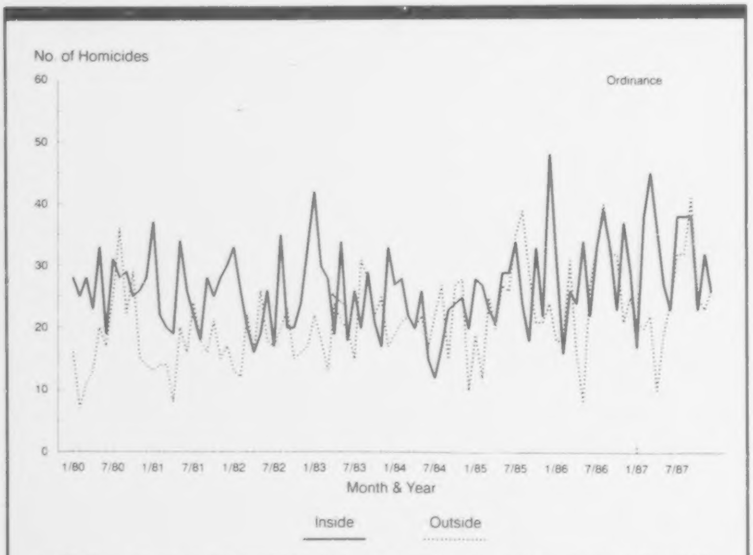

FIGURE 2—Number of Homicides in Detroit by Month and Location of Homicide, 1980–87

Enforcement of the Ordinance

The research design assumes that the ordinance resulted in the mandated change in jail sentences for illegally concealing a pistol or carrying a firearm. We determined, however, that the law did not result in any such change. The records of the 1,020 persons charged under the ordinance whose cases were processed as of early 1989: only 22 defendants (2.2 percent) were sentenced to jail and eight others (0.8 percent) were given credit for jail time served while waiting disposition of their case. These findings were confirmed informally through our interviews with key criminal justice personnel. Two years after the ordinance was passed, everyone we interviewed who was in a position to observe sentencing practices agreed that extremely few of those who were arrested for carrying concealed weapons ultimately received the mandatory sentences stipulated in the ordinance.

The administrative context in which the ordinance is enforced explains why it failed to increase the certainty or severity of sentences. Since long before the ordinance was passed, officials at the Wayne County Jail had been under court order to reduce overcrowding and improve living conditions. This order had transformed jail space into a scarce and expensive commodity. Judges had to choose who among the misdemeanants, ordinance violators, and pretrial felony defendants to incarcerate in the 1,676-bed county jail. Our informants agreed unanimously that it was impossible to enforce the mandatory sentencing provisions of the ordinance because not enough jail space was available. Given a choice between incarcerating violators of the gun ordinance versus pretrial felony defendants or other offenders who had committed more serious offenses such as violent assault or robbery, ways were found to avoid the mandatory sentencing provision of the gun ordinance.

TABLE 1—Summary of Interrupted Time-series Analysis of the Effect of a Gun Ordinance on the Incidence of Homicide, Detroit, Michigan, 1980 to 1987

Homicide Series	Estimate of Intervention Effect	Mean Number of Homicides[a]	Percentage Change in Series[b]	P-value for Test of H_0 = no change
Inside	5.798	25.95	22	.0058
Outside	2.086	20.83	10	.4179
Nonfirearm	2.524	15.56	16	.0703
Firearm	4.109	31.42	13	.2380

[a] The mean number of homicides per month in the preintervention series adjusted, where appropriate, for autocorrelation.

[b] The intervention effect as a percent of the mean of the preintervention series.

Discussion

Unquestionably, homicide is an enormous public health problem in the United States.[12] Viewed strictly from a health perspective, the numbers of deaths and injuries from assaultive violence have certainly reached epidemic proportions in some American cities. Public health authorities have rightly called for increased efforts to address this urgent health problem by whatever means are determined to be both effective and societally acceptable. The descriptive epidemiology of homicide is clear on three salient points: 1) many homicide victims are killed by friends or acquaintances; 2) many homicides occur as a result of an argument, frequently over relatively trivial matters; and 3) approximately six out of 10 homicides are committed with firearms, usually handguns.[13] On the basis of these observations, several investigators have hypothesized that many homicides could be prevented if ready access to firearms were reduced, since many violent altercations occurring among family, friends, and acquaintances would then necessarily be resolved by less deadly means.[14-16]

The Detroit gun ordinance evaluated here is one approach to decreasing the accessibility of firearms. Theoretically, by reducing the number of persons carrying firearms on the streets of Detroit, some violent public altercations that might have resulted in fatal gunshot injuries would result in other nonfatal outcomes. Our time-series findings are consistent with a model in which the ordinance had a relatively small preventive effect on the incidence of homicides in Detroit. More importantly, the effect is found precisely where we predicted we would find it, based on our hypothesis that the ordinance would affect the incidence of firearm homicides and homicides occurring outside but not the incidence of other types of homicide.

We also found that the Detroit ordinance requiring mandatory sentences was almost completely unenforced. If the ordinance was never enforced, how can any preventive effect, small or otherwise, be ascribed to it? During the period from November 1986 through January 1987, media coverage of the passage of the ordinance was extensive and public debate about the virtues of the ordinance was heated. In several articles, local officials pledged to enforce the law strictly and to provide jail space as necessary. This publicity may have been responsible for the preventive effect that we found associated with the promulgation of the ordinance. The lack of enforcement—and the coincident ending of the publicity surrounding the law's passage after January 1987—may account for the relatively small nature of the observed effect.

Little previous research has been done on the effectiveness of laws that require mandatory sentences for carrying firearms illegally. To date, the most extensive studies of this issue dealt with the Massachusetts Bartley-Fox Law which required a mandatory minimum one-year sentence for carrying a firearm illegally. Investigators reported that the Bartley-Fox law reduced the number of robberies, assaults, and homicides committed with a firearm.[5,17] A more in-depth study of assaults, however, revealed that, whereas assaults with firearms decreased, assaults without firearms increased. This raises the question of whether a reduction in firearm-carrying results in a displacement rather than a reduction in interpersonal violence. In the case of the Bartley-Fox law, the apparent increase in nonfirearm assaults appears to have been due to increased reporting of more minor assaults after the law was implemented.[17] From a public health perspective, however, this issue is moot: a displacement of interpersonal violence from altercations in which firearms are used to altercations in which firearms are not used would most likely have the desired effect of preventing many homicides.

In addition to the intuitive appeal of reducing homicides in public by reducing the public carrying of firearms, ordinances such as this one in Detroit are worthy of close evaluation for several other reasons. First, the law imposes relatively few costs on individual citizens who own and use firearms for purposes such as hunting or protection. The ordinance does not prevent a person from keeping a firearm in the home. In addition, under the ordinance a firearm may be legally transferred from place to place (e.g., from the home to a firing range or out to a hunting area) as long as it "is unloaded in both barrel and magazine and is taken down, enclosed in a case, carried in the luggage compartment of the vehicle or inaccessible from the interior of the vehicle." The ordinance only prohibits carrying a concealed, loaded pistol on one's person, or any

APPENDIX

Noise Model:	Intervention Model:
Inside Homicides	
ARIMA(0,0,0) $a_t = Y_t - \Theta_0$ $\hat{\Theta}_0 = 25.95$ s.e. = .7429 Box-Ljung Q = 24 with 25 degrees of freedom	$Y_t^* = a_t + \omega_0 (I_t)$ $I_t = 1$ for months 1–84 $I_t = 0$ for months 85–96 $\hat{\omega} = 5.798$ s.e. = 2.1011
Outside Homicides	
$\text{ARIMA}(1,0,0)(1,0,0)_{12}$ $a_t = (1 - \phi_1 B)(1 - \phi_{12} B^{12}) Y_t - \Theta_0$ $\hat{\Theta}_0 = 20.83$ s.e. = 1.3496 $\hat{\phi}_1 = .3220$ s.e. = .0970 $\hat{\phi}_{12} = .3529$ s.e. = .0957 Box-Ljung Q = 19 with 23 degrees of freedom	$Y_t^* = a_t + \omega_0 (I_t)$ $I_t = 0$ for months 1–84 $I_t = 1$ for months 85–96 $\hat{\omega} = 2.086$ s.e. = 2.5771
Nonfirearm Homicides	
ARIMA(0,0,0) $a_t = Y_t - \Theta_0$ $\hat{\Theta}_0 = 15.56$ s.e. = .4920 Box-Ljung Q = 15 with 25 degrees of freedom	$Y_t^* = a_t + \omega_0 (I_t)$ $I_t = 0$ for months 1–84 $I_t = 1$ for months 85–96 $\hat{\omega} = 2.524$ s.e. = 1.3918
Firearm Homicides	
$\text{ARIMA}(1,0,0)(1,0,0)_{12}$ $a_t = (1 - \phi_1 B)(1 - \phi_{12} B^{12}) Y_t - \Theta_0$ $\hat{\Theta}_0 = 31.42$ s.e. = 1.54[illegible] $\hat{\phi}_1 = .2791$ s.e. = .1016 $\hat{\phi}_{12} = .2301$ s.e. = .1115 Box-Ljung Q = 19 with 23 degrees of freedom	$Y_t^* = a_t + \omega_0 (I_t)$ $I_t = 0$ for months 1–84 $I_t = 1$ for months 85–96 $\hat{\omega} = 4.109$ s.e. = 3.4885

loaded firearm in a car in such a manner that it could readily be retrieved.

Second, the notion that local governments have the right to determine who may legally carry loaded firearms in their jurisdictions appears to be relatively well accepted and not subject to much of the controversy surrounding "gun control." Indeed, the Detroit gun ordinance may be considered innovative only because of the inclusion of mandatory sentencing provisions. If such laws are generally considered to represent a reasonable function of local government, then—should scientific study demonstrate their utility in preventing homicide—ordinances like the one in Detroit might be adopted in other jurisdictions with relatively little controversy.

In contrast to the minor costs imposed by the ordinance on individual citizens, however, the general costs associated with arrest, adjudication, and mandatory incarceration of violators of the ordinance are *not* trivial. As noted in our evaluation of the enforcement of the Detroit ordinance, jail time is frequently both a scarce and expensive resource, and mandatory sentence laws, if enforced, would certainly increase the numbers of persons in jail, at least in the short term.

Given the dearth of similar research, the relatively small magnitude of the effect observed in Detroit, and the fact that the Detroit ordinance was not enforced, our findings must not be considered definitive. Reasonable caution requires that the study be replicated in a setting in which a similar ordinance is effectively and visibly enforced. Despite the theoretical and practical appeal of the Detroit gun ordinance, assessing the effectiveness of such ordinances in preventing homicide is imperative before public health officials advocate their wider adoption by other localities. In addition, the feasibility of proper, visible, and ongoing enforcement of such laws should also be assessed. Merely passing an ordinance to decrease the carrying of loaded firearms in public is clearly not adequate. Moreover, passing but not enforcing such ordinances may have unanticipated negative consequences. The inevitable ineffectiveness of an unenforced ordinance might lead policy makers to the inappropriate conclusion that such ordinances are *inherently* ineffective. If laws are passed to curtail the number of people carrying loaded firearms in public, they must be enforced. If such laws are passed *and* enforced, then the public health community should miss no opportunity to evaluate the effect of these laws on the rate of injury and death from firearms violence in public. □

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The statistical analysis presented herein was independently reviewed and confirmed by Daniel J. Schnell, PhD, Mathematical Statistician, Division of STD/HIV Prevention, Center for Prevention Services, Centers for Disease Control.

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Fellowship Fund at UCLA to Benefit American Indian Students

The George Blue Spruce, Jr., Fellowship Fund for American Indian Students was recently established at the University of California at Los Angeles, School of Public Health. Named in recognition of George Blue Spruce, Jr., of Phoenix, Arizona, the fellowship will provide for scholarships and loans to American Indian students preparing for careers in the field of public health. The fellowship is administered by the Health Careers Opportunity Program at the school, under direction of a committee composed of American Indian alumni from the UCLA School of Public Health.

Dr. Blue Spruce, who comes from the San Juan/Laguna Pueblo Tribes of New Mexico, was the first American Indian to graduate as a dentist in the United States. His doctorate of dental surgery is from Creighton University. Blue Spruce's career in health service includes work as a US Navy dental officer, clinical assignments on Indian reservations, assistant surgeon general in the US Public Health Service, director of the Office of Native American Programs in the US Office of Human Development, and consultant in dental health for the World Health Organization, In addition, he has chaired the Intra-Departmental Council of Indian Affairs in Washington, DC and was director of the Phoenix Area Indian Health Services for the Department of Health Services.

To date, approximately 30 American Indians have completed advanced degree programs at the UCLA SPH. The George Blue Spruce, Jr., Fellowship Fund is the third fund established at the school to support the recruitment and retention of minority students. In 1987, the Richard Chavez Scholarship Fund was established for Hispanic students, and in 1990 the Betty Smith Williams Scholarship Fund was established for African-American students.

For further information and donations to the fellowship, contact: Health Careers Opportunity Program, UCLA School of Public Health, 41-240, Los Angeles, CA 90024-1772.

Cervical Papillomavirus Infection and Cervical Dysplasia in Hispanic, Native American, and Non-Hispanic White Women in New Mexico

ABSTRACT

Background: Human papillomavirus infections of the cervix are found with varying frequencies in different populations worldwide, and have been associated with cervical cytologic abnormalities.

Methods: We studied 1,603 randomly selected Hispanic, Native American, and non-Hispanic White women in New Mexico to determine the prevalence of cervical HPV infection in these ethnic groups, and its association with Pap smear abnormalities, using a new commercial dot-blot hybridization assay.

Results: Nine percent of all women screened had evidence of cervical HPV infection (13.7% of non-Hispanic White women, 9.7% of Hispanics, and 6.6% of Native American women). Prevalence was higher in younger women ages 14–19 years than in older age groups. Over half of women with cervical HPV infection (n = 145) had normal Pap smears. The proportion of infected women increased among those with more advanced cytopathologic abnormalities; 5.6 percent with normal Pap smears had cervical HPV vs 66.7 percent with moderate-severe dysplasia.

Conclusions: Cervical HPV infection is common among New Mexico clinic attendees, varies in prevalence among the three major ethnic groups, and is strongly associated with cervical cytopathologic abnormalities. (*Am J Public Health* 1991;81:582–586)

Thomas M. Becker, MD, Cosette M. Wheeler, PhD, Nancy S. McGough, RN, Scott W. Jordan, MD, Maxine Dorin, MD, and Jill Miller, MD

Introduction

Human papillomavirus (HPV) infections of the genital tract are among the most common sexually transmitted viral infections in the United States.[1-3] Recent estimates have suggested that as many as 12 million adults in this country may be infected with genital HPV.[3] In addition to causing genital warts, HPV infections may also result in clinically inapparent lesions of the cervix. Cervical HPV infection can be diagnosed by certain morphologic features on colposcopy,[4] by cytopathologic changes seen in Papanicolaou (Pap) smears,[5] through detection of HPV antigen by immunoperoxidase stains,[6] or by identification of HPV DNA sequences in hybridization studies.[7,8] Hybridization techniques, considered more sensitive and specific than most other types of HPV tests,[9] have shown the presence of HPV genomes in 6–30 percent of women tested for cervical abnormalities from various populations worldwide.[10-15] The use of recently developed polymerase chain reaction-based HPV diagnostic tests suggests that cervical HPV prevalence may be even higher than has been determined by hybridization methods.[16,17]

New Mexico's population is comprised primarily of three major ethnic groups—Hispanics, Native Americans, and non-Hispanic Whites. In prior reports, we have shown variation in the incidence and prevalence of many infectious and chronic diseases among these ethnic groups,[18-22] including invasive and pre-invasive cervical cancer.[23] As part of an ongoing cohort study to investigate the natural history of cervical HPV infection, we studied Hispanic and non-Hispanic White women attending the University of New Mexico (UNM) gynecology clinics, and Native American women attending local Indian Health Service (IHS) hospital clinics to determine ethnic differences in the prevalence of cervical HPV infection using a newly developed hybridization test, and the relationship of cervical HPV infection to cervical cytologic abnormalities in these ethnic populations.

Methods

Study Sample

Subjects were women who attended the UNM gynecology clinics and the Albuquerque IHS hospital-affiliated clinics, June 1988–September 1989. On randomly selected days of the week, providers screened all women on whom a pelvic examination was performed for any reason. This sampling scheme was consistent among clinics and sites and all five clinic days (Monday–Friday) were equally likely to be selected. Patients who underwent a pelvic examination for any reason were included in this survey. The UNM clinic attendees were comprised primarily of Hispanic and non-Hispanic White women; few women of other ethnic or racial groups attend these clinics. At the UNM clinics, Hispanic ethnicity is either

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Address reprint requests to Thomas M. Becker, MD, Assistant Professor, Department of Medicine, Cancer Center, School of Medicine, University of New Mexico, Albuquerque, NM 87131. Dr. Wheeler is Research Assistant Professor, Department of Cell Biology; Ms. McGough is Research Nurse, Department of OB/GYN; Dr. Jordan is Professor, Department of Pathology; Dr. Dorin is Assistant Professor, Department of OB/GYN; Dr. Miller is Clinical Adjunct Professor, Department of Family Medicine; all at UNM School of Medicine. This paper, submitted to the Journal April 17, 1990, was revised and accepted for publication January 8, 1991.

determined by self-identification at the time of clinic enrollment or is assigned by clinic staff on the basis of Spanish surname. To validate the UNM clinics' designation of Hispanic and non-Hispanic White ethnicity, we interviewed a 15 percent sample of the UNM clinic group studied. We found 91 percent concordance between the self-reporting of Hispanic or non-Hispanic White ethnicity compared to the combined self-identified/clinic staff assignment of ethnic status. The Indian Health Service hospital clinics serve primarily Pueblo and Navajo women. No attempt was made to validate Native American ethnicity. We did not include in our report those Black women who were screened at the UNM clinics for cervical HPV (n = 7). Informed consent was required at three of the four clinics affiliated with the UNM gynecology department. Only two refusals to participate in the screening survey were recorded among all clinic patients; both of these patients were of Hispanic ethnicity.

UNM clinics serve primarily women who are economically disadvantaged. The IHS-affiliated clinics provide services free of charge to eligible Native American patients. Nonetheless, Native Americans in New Mexico are an economically disadvantaged group: approximately 40 percent of the state's Native American population lived below the poverty level in 1980.[24]

Cytopathologic Diagnoses

Pap smears from the UNM clinics were all read by the same cytopathologist (SWJ); cytologic smears from the IHS hospital were initially read by pathologists from Path Labs, El Paso, Texas. We found 98 percent concordance between these two laboratories for diagnoses of a sample of normal cervical smears (n = 59). The UNM cytopathologist (SWJ) reviewed all abnormal Pap smears which were initially reported by Path Labs for the IHS clinic attendees. To maintain consistency in cytopathologic diagnoses for this report, we used the UNM cytopathologist's diagnoses for abnormal Pap smears in our analysis. Cytopathologic changes associated with HPV infection (such as koilocytosis) were coded as atypias for this report. All diagnoses corresponded to the World Health Organization's criteria for cervical cellular abnormalities.[25] During the period of this survey, the newly developed Bethesda Classification System for cervical cytology had not been implemented in either pathology laboratory.[26]

HPV Testing

We used a recently developed dot hybridization screening test, ViraPap® (Life Technologies, Inc, Gaithersburg, MD 20857) to test for cervical HPV infection. This test is designed to identify genital HPV types 6, 11, 16, 18, 31, 33, and 35—the types most commonly associated with genital tract infection. The test includes sampling the endo- and exocervix with a dacron-tipped plastic swab for exfoliated cells. HPV samples were collected after the Pap smears were collected. Dacron swabs were placed immediately in tubes of ViraPap™ collection media. The tubes were frozen at −20°C until testing in batch at the HPV laboratory, University of New Mexico School of Medicine. All positive HPV ViraPap tests underwent typing using the ViraType® test (Life Technologies, Gaithersburg, MD 20857). This dot hybridization test is able to separate HPV genome into the three categories, 6/11, 16/18, and 31/33/35, but does not determine which of the virus types is present within a given category. All borderline test results by ViraPap were further tested by ViraType and considered negative only if they were negative by the ViraType test.

The HPV laboratory independently validated the sensitivity and specificity of ViraPap and ViraType using Southern Blot hybridization techniques as the "gold standard".[16,17] For 147 specimens, the sensitivity of the ViraPap assay was 100 percent and specificity 96 percent. For 63 specimens, the sensitivity of the ViraType assay was 100 percent and specificity 91 percent.

Data Analysis

Data were analyzed using standard packages from SAS.[27] We examined the age-adjusted prevalence of cervical HPV infection for Hispanic, Native American, and non-Hispanic White women between the ages of 15 to 44 years using the combined population distribution for these study patients in this age range as our reference population. We age-adjusted these figures using the direct method of standardization.[28]

Results

We studied 535 Hispanic, 322 non-Hispanic White, and 746 Native American women, a total of 1,603 women. The mean ages of the participants were 25.0 years for the Hispanic women (range 15–43 years), 25.5 years for non-Hispanic Whites (range 14–50 years), and 29.5 years for Native American women (range 14–78 years). The prevalence of cervical HPV infection in this sample was higher for non-Hispanic White women (13.7 percent) than for Hispanic (9.7 percent), or Native American women (6.6 percent). The age-adjusted prevalence of cervical HPV infection was 13.1, 9.3, and 7.7 percent for non-Hispanic White, Hispanic, and Native American women, respectively, between the ages of 15 to 44 years. Among all women with cervical HPV infection, Native American women had a higher proportion of HPV 6/11 compared to non-Indian women (Table 1).

Younger women were more likely to have cervical HPV than were women in older age groups (Table 2). Comparable age-related decreases were observed in each of the three ethnic populations which we studied (data not shown). For all women combined, 6.7 percent of patients screened had slight, moderate, or severe dysplasia. American Indian women had a higher proportion of Pap smears showing atypias, compared to Hispanic and non-

TABLE 1—Cervical HPV Test Results by Ethnic Group and HPV Type, Cervical HPV Infection Survey (n = 1603)

		HPV Types			
Ethnic Group	HPV Positive No. (%)	6/11 No. (%)	16/18 No. (%)	31/33/35 No. (%)	Mixed[b] No. (%)
Hispanic n = 535	52 (9.7)	3 (5.8)[a]	32 (61.5)	16 (30.8)	1 (1.9)
non-Hispanic White n = 322	44 (13.7)	1 (2.3)	15 (34.1)	26 (59.1)	2 (4.5)
American Indian n = 746	49 (6.6)	6 (12.2)	18 (36.7)	23 (46.9)	2 (4.2)

a) Percent of all HPV-positive specimens within that ethnic population.

b) Mixed virus types include combinations of 6/11 and 31/33/35, (n = 1), and 16/18 and 31/33/35 (n = 2), and 6/11 and 16/18 (n = 2).

TABLE 2—Cervical HPV Test Results by Age Group for all Ethnic Groups Combined, Cervical HPV Infection Survey (n = 1603)

HPV Results	Age Group (years) 14–19 No. (%)	20–29 No. (%)	30–39 No. (%)	40–49 No. (%)	50–59 No. (%)	60+ No. (%)	Unknown No. (%)
Negative	159 (86.9)	811 (89.7)	378 (93.8)	69 (98.6)	20 (100)	13 (100)	6 (75.0)
Positive	24 (13.1)	93 (10.3)	25 (6.2)	1 (1.4)	0	0	2 (25.0)
Total	183	904	403	70	20	13	8

TABLE 3—Cytopathologic Diagnoses for Hispanic, Non-Hispanic White, and American Indian Participants, Cervical HPV Infection Survey

	Negative No. (%)	Atypia No. (%)	Slight Dysplasia No. (%)	Moderate-Severe Dysplasia No. (%)
Hispanic n = 535	499 (93.3)	4 (.8)	25 (4.7)	7 (1.3)
non-Hispanic White n = 322	284 (88.2)	3 (.9)	27 (8.4)	8 (2.5)
American Indian n = 746	664 (89.0)	45 (6.0)	25 (3.4)	12 (1.6)
Total n = 1603	1447 (90.3)	52 (3.2)	77 (4.8)	27 (1.7)

Hispanic white women (Table 3). Table 4 shows that a large proportion of the 145 women with evidence of cervical HPV infection had normal Pap smears (55.9 percent), but that the prevalence of HPV increased sharply with the degree of dysplasia found in the Pap smear.

Discussion

Our data show ethnic differences in the prevalence of cervical HPV infection, as well as ethnic differences in types of HPV present on the cervix at the time of the screening examination, and substantiate prior published observations that HPV is more commonly found in more advanced cervical lesions.[3,12] Unlike most prior reports, however, our survey showed that HPV types 31, 33, and 35 were often present in association with cytologic abnormalities.

The prevalence of cervical HPV infection in different populations worldwide varies substantially, depending on the diagnostic test employed for HPV detection. Using cytologic criteria, Meisels and Morin[5] reported that 1.7 percent of 234,715 women who participated in a screening program in Quebec, Canada, between 1975–1979 had cytologic evidence of cervical HPV infection. In other cytologic surveys, prevalence estimates ranged from 0.9 percent to 16 percent.[1,29–33] Among these reported series, the prevalence has been higher in sexually transmitted disease clinics[29,32,33] than in other clinic populations. Research which has employed DNA hybridization techniques found evidence of cervical HPV DNA to vary from 6 to 28 percent of women tested,[10–15] with several studies reporting a prevalence of about 10 percent.[11–15,29,33] Using Southern blot hybridization for evidence of HPV 6, 11, 16, or 18 DNA, deVilliers, *et al*,[12] found the highest prevalence of HPV DNA was from cervical specimens from women less than 25 years of age (14 percent), and the lowest prevalence among women age 56 or older (4 percent). Similarly, Kiviat, *et al*, reported a higher prevalence of cervical HPV DNA in women 16–24 years old who attended a sexually transmitted disease clinic or a student health clinic in Seattle, Washington compared with older age groups of women at both clinics.[33]

HPV types 16, 18, and 31 have been implicated as having a more aggressive role in malignant transformation of cervical cells.[34–38] Our screening data cannot address the issue of malignant transformation; however, if published observations on the malignant potential of HPV 16, 18, and 31 are accurate,[11,36] many of the clinic attendees we studied in New Mexico are at high risk for development of cervical cancer. Nonetheless, current epidemiologic evidence is too weak to accurately assess the relationship of specific HPV types in genital cancer etiology: carefully designed prospective studies are necessary which consider other potential risk factors in the development of disease. Ethical considerations, however, cannot allow clinicians to follow patients to the end-point of cervical cancer without earlier intervention; thus, the appropriate studies to determine the relationship between HPV infection and cervical cancer will probably never be carried out. Population-based screening for HPV (by type), together with examination of cervical cancer specimens for type-specific HPV genome from cases which occur in those same populations, may shed additional light on the relationship of specific HPV types in cervical cancer.

Although several case-control studies and one cohort study have indicated that the presence of HPV in cervical tissue is a strong risk factor for cervical dysplasia

TABLE 4—Cytopathologic and HPV Test Results by HPV Type, all Ethnic Groups Combined, Cervical HPV Infection Survey (n = 1603)

Pap Smear Diagnosis	HPV (+) No. (%)	HPV type 6/11[a] No. (%)	16/18 No. (%)	31/33/35 No. (%)	Mixed No. (%)
Negative n = 1447	81 (5.6)	8 (9.9)	34 (42.0)	37 (45.7)	2 (2.5)
Atypia n = 52	11 (21.2)	0 (0)	7 (63.6)	4 (36.4)	0 (0)
Slight n = 77	35 (45.5)	2 (5.7)	13 (37.1)	17 (48.6)	3 (8.6)
Mod-Severe n = 27	18 (66.7)	0 (0)	11 (61.1)	7 (38.9)	0 (0)

a) % of all specimens which showed evidence of HPV within the Pap smear diagnostic category.

or cervical cancer,[2,10,12,14,38,39] data on genital cancer incidence in New Mexico's Hispanics, Native Americans, and non-Hispanic Whites when compared to our survey findings appear to provide evidence that conflicts with these studies. The highest cervical cancer rates in New Mexico occur among Native American and Hispanic women, with rates for non-Hispanic White women which are comparable to those rates for whites nationwide[23]—the exact opposite of our survey of HPV prevalence. Although such ecological correlations are weak and do not consider numerous confounding factors in the infection-disease relationship, they suggest that we examine the epidemiology of cervical disease as related to HPV infection and to other relevant behavioral, infectious, and dietary factors more closely.

Our study has limitations that must be addressed. The HPV laboratory test we used, while demonstrating high sensitivity and specificity compared to Southern Blot hybridization for certain HPV types, is unable to determine the presence of genital HPV types other than those included in the test kit. The New Mexico clinic populations may thus contain an even greater proportion of women infected with other genital HPV types which have been identified in other populations.[40–42] Furthermore, implementation of highly sensitive polymerase chain reaction-based HPV diagnostic testing in other populations suggests that cervical HPV prevalence is considerably higher than can be determined by using hybridization methods.[16,17] Thus, it is likely that cervical HPV prevalence in our clinic populations may be higher as assessed through polymerase chain reaction techniques than we report using the commercial dot-blot assay. In addition, we have examined single specimens for the presence of cervical HPV infection; serial sampling of the same patients may result in the determination of an even greater number of women with cervical HPV infection. Poor technique in the collection of Pap smears or HPV samples could also influence our results; however, all clinicians who collected these specimens were instructed in proper sample collection techniques. Despite these potential sources of bias in our study, we have demonstrated that cervical HPV infection is common among university-affiliated and Indian Health Service hospital-affiliated clinic attendees in New Mexico; prospective studies are ongoing to evaluate the relationship of cervical HPV infection to the development of dysplastic lesions in these populations. □

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New Report from Bureau of Health Professions

The Health Resources and Services Administration's Bureau of Health Professions (BHPr) recently announced publication of the summary report of its *Workshop on Modeling and Forecasting for the 1990s* (DHHS Pub. No. HRS-P-OD-90-5). This invitational workshop was sponsored by BHPr's Office of Data Analysis and Management. It focused on the desirability of utilizing a common framework of economic and health sector assumptions for all Bureau modeling and forecasting efforts.

The report summarizes the formal presentations and structured discussions of the more than 50 health services researchers from federal and state government, universities, and private research firms at the January 1990 workshop. Appendices include the workshop keynote address, agenda, schematics of BHPr models, selected references, and participant feedback.

Single copies of the report may be obtained by writing the Office of Data Analysis and Management, Howard V. Stambler, Director, BHPr, Rm. 8-47 Parklawn Bldg., 5600 Fishers Lane, Rockville, MD 20857; (301) 443-6936.

Urolithiasis in Tennessee: An Occupational Window into a Regional Problem

ABSTRACT

Background: Urinary tract stones (stones) are believed to be unusually common in the southeastern United States but neither the incidence of nor the risk factors for stones are known.

Methods: In three well-defined occupational populations in eastern Tennessee, we assessed the prevalence, incidence, and cumulative incidence of stones and measured biochemical risk factors for lithogenesis.

Results: The age-adjusted prevalence of stones was 18.5 percent in Tennessee compared to 7.7 percent among White males in US NHANES (prevalence ratio 2.4, 95% CI 1.7, 3.3). The cumulative incidence (risk) was 27.8 percent by age 65, higher than in any other reported population. Risk factors were age, a family history, and urinary saturation with calcium-oxalate (COAX). Persons with a positive family history and the highest measured CAOX index had a predicted lifetime risk of 88.8 percent. Biochemical factors affecting lithogenesis were hypercalcuria, hyperoxaluria, and low urine volume.

Conclusion: Future research should characterize the geographic boundaries of a southeastern "stone-belt" and explore genetic and dietary hypotheses that might explain it. (*Am J Public Health* 1991;81:587–591)

Michael J. Thun, MD, MS, and Susan Schober, PhD

Introduction

Kidney stones, an important source of morbidity, suffering, and medical expense, have received little epidemiologic attention. Three studies suggest an endemic area of stones in the Southeastern United States, but these rely upon hospital discharge data rather than incidence.[1–3] The absence of population-based incidence data in the US and the paucity of etiologic studies impede the identification of potentially modifiable risk factors such as diet and climate.[4,5]

The present study began as an investigation of reported urinary tract disorders among uranium workers in eastern Tennessee. Although no occupational association was found,[6] the study provided an opportunity to assess stone occurrence and risk factors for lithogenesis in a high-risk population. Our data support the concept of a regional "stone belt" and suggest clues as to its etiology.

Methods

Subjects

The uranium plant processed depleted and natural uranium and thorium from 1957 to 1970, and enriched uranium after 1970. Workers eligible for the study represent a gradient of exposure to unenriched uranium, compounds for which nephrotoxicity rather than radiation limits exposure.[7–9] The target population included: 1) highly exposed active, retired, and disabled production workers with 20 years employment before 1986; 2) a low exposure group of senior office and laboratory workers; 3) minimally exposed guards, ages 40–65, employed recently at the uranium plant; and 4) unexposed dairy workers from a nearby town. The dairy workers belong to the same trade union and local community. They produce tinned evaporated milk but have no special access to discounted milk or cheese that might affect stone risk.

Initially, a pilot questionnaire was administered to senior uranium workers and to guards. Later, a more extensive questionnaire and medical study were offered to the entire target population. Both questionnaires asked identical core questions about age and stones:

- "Have you ever had a kidney stone, including gravel in your urine?"
- "How many times have you had this?"
- "When did you first have this?"
- "Where were you treated?"

Workers interviewed in person were asked to sign a medical release form. The medical questionnaire also asked about a family history of stones, treatment for urinary tract infections, length of residence in the area, diabetes, hypertension, urinary instrumentation, prostatic disease, analgesics, smoking, "moonshine," and occupational exposure to heavy metals and solvents.

Data Collection

This activity included height, weight, seated blood pressure, serum for total calcium, phosphorus, uric acid and creatinine, sterile urine for quantitative urine culture and microscopy, and urine samples for chemistry, beta-2-microglobulin,

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From the National Institute for Occupational Safety and Health, Cincinnati, Ohio. Address reprint requests to Michael J. Thun, MD, MS, Department of Epidemiology and Statistics, American Cancer Society, 1599 Clifton Road, NE, Atlanta, GA 30329. Dr. Schober is currently with the National Institute on Drug Abuse, Rockville, MD. This paper, submitted to the Journal March 29, 1990, was revised and accepted for publication October 19, 1990.

TABLE 1—Participation in the Questionnaire and Medical Study

Employment Category	Number Eligible	Participation Core Questionnaire N (%)	 Full Medical Study (%)*
Uranium Exposed			
Active			
Hourly	58	58 (100)	50 (89)
Salaried	36	36 (100)	19 (59)
Retired			
Hourly	6	5 (83)	2 (40)
Salaried	10	10 (100)	2 (20)
Disabled	6	6 (100)	3 (50)
Referents			
Guards	43	43 (100)	10 (24)
Dairy Workers	51	50 (98)	37 (79)
Total	210	208 (99)	123 (62)

*Participation in medical study based upon the number (199 total) of workers present locally when testing performed.

retinol binding protein, and enzymes.[6] Urine was collected over 24 hours to measure the "stone risk profile," including eight promoters of lithogenesis (calcium, oxalate, sodium, uric acid, sulfate, phosphorus, and high or low pH), and three inhibitors (citrate, magnesium, and urine volume).[10,11] During collection, participants maintained their usual diet, fluid intake, and activity. Urine pH was measured immediately upon return of the sample; volume was determined by dilution of a known quantity of lithium. Aliquots were shipped in dry ice for analysis.[6]

Supersaturation with CAOX, dicalcium phosphate, monosodium urate, and urate was computed from measured values using a computer algorithm that considers the complex physiochemical interaction between multiple factors.[12] Supersaturation scores are expressed relative to the group mean of 176 persons without stones.[12]

Statistical Analyses

Age-specific prevalence reflects the number of workers ever having had a stone divided by the number in the corresponding age bracket, based upon age at interview. Age-adjusted prevalence was compared to White males in NHANES II[13] using maximum likelihood estimation.[14] Age-adjusted incidence reflects the age at first stone. Person years at risk (PYAR) were computed using life-table methods. Workers contribute PYAR from birth until they develop a stone or complete the interview, whichever is first. Age-adjusted incidence in Tennessee was compared to published data from the Mayo clinic,[5] using maximum likelihood estimation;[14] 95% confidence intervals used the test-based method.[14] Cumulative incidence (based on age at first stone) used the Kaplan-Meier method.[15] Factors associated with stones were identified using Student's t-test (for continuous variables) and Chi-square (for proportions). Associations were further tested using unconditional logistic regression. In this analysis, cases are persons ever reporting stones. Terms considered in the regression are age and all factors in Table 4 associated with stones at $p < 0.2$. The effect of each risk factor was assessed separately and in the presence of others. Age is retained in the final model because of the known association of stones and age. Other factors are retained if statistically significant ($p < 0.05$) or if inclusion of the factor substantially influences the coefficient for other terms. Second degree interaction was tested using stratified analyses; interaction terms are not included in the logistic model because none of these substantially modify the main effect terms.

Results

Table 1 shows participation rates by employment category. Overall, 208 eligible workers (99 percent) completed the core questionnaire. Of 199 present locally during the medical testing, 136 (68 percent) completed the medical questionnaire and 123 (62 percent) provided blood and 24-hour urine samples (full medical study). Participation was highest among active hourly workers released from duty for testing. Guards, retired, and salaried workers were either off-site, not released from duty, or had lower commitment to the study. All participants were male, 206 were White, with an average age of 52.1 and range of 38–72 years.

Medical Record Review

Forty of 208 workers reported stones (crude prevalence 19.2 percent). Of these, 37 were treated by civilian physicians, 30 (all those interviewed in person) provided medical release forms, 28 received records from the treating facility, and 25 had a stone confirmed. Stones could not be confirmed when the treating physician had retired or relocated. Confirmed stones accounted for 34 hospitalizations, 42 abdominal X-rays, five surgical extractions, and five lithotripsies. Chemical analyses of stones were recorded on 13 persons with recurrent stones; calcium oxalate was present in 12 (92 percent). Nineteen persons reported multiple stones; five of these had from five to 20, and one had 32.

Prevalence

Table 2 presents the prevalence of stones by age and occupational group. At all ages, the Tennessee workers reported more frequent stones than did White males in NHANES.[13] The age-adjusted prevalence in the Tennessee workers was 18.5 percent vs 7.7 percent in NHANES (prevalence ratio 2.4, 95% CI = 1.7, 3.3). Stone prevalence was similar in all occupational groups except guards. Prevalence in guards was 7 percent, significantly lower than that of other workers ($p = 0.02$) and similar to NHANES.[13]

Incidence and Cumulative Incidence

The age-adjusted incidence of first stones was 4.21 per 1,000 person-years, compared to 1.03 per 1,000 in Minnesota males[5] (RR = 4.1, 95% CI = 2.9, 6.0). The cumulative incidence (risk) of stone by age 60 was 23.7 percent, and by age 65 was 27.8 percent.

Characteristics Associated with Stones

Table 3 shows selected demographic and clinical attributes of persons with and without stones. Stone-formers in our study weighed more, used moonshine liquor less, reported a more frequent family history of stones and of male relatives treated for urinary infections, and were more likely to be treated for such infections themselves. Laboratory findings associated with stones are a significantly higher serum and urine calcium, and urinary supersaturation with CAOX and brushite. Findings not associated with stones in our study but of interest in the literature are serum and urine uric acid,[16] current cigarette smoking,* and analgesic use.[17]

Table 4 presents the results of unconditional logistic regression. The two factors most strongly associated with stones were the index for CAOX supersaturation in a current urine sample, and a family history of stones. The logistic model (Table 4) categorizes the CAOX index into tertiles. In Figure 1, the CAOX index is presented as a continuous variable. Figure 1 illustrates the probability of stones, as predicted by the model, in persons with and without a family history over the range of the CAOX observed in our study. Males with a family history and the highest measured CAOX index have a predicted lifetime probability of stones of 88.8 percent. Using multiple linear regression, we determined that the most important determinants of the CAOX were high urinary calcium and oxalate and low urine volume. Together, these terms explained 91.8 percent of the variance, with the individual contribution of calcium, urine volume, and oxalate being 49.5 percent, 34.1 percent, and 8.2 percent, respectively. Six other stone risk factors used to compute the CAOX index were statistically significant at the $p < 0.001$ level, but together these added only 6.7 percent to the explained variance.

TABLE 2—Age-Specific Prevalence of Stones: Tennessee Workers vs US

	Number (Percent) with Stones			
	Tennessee (N = 208)		US NHANES[18] (N = 4331)	
Age (years)	N	(%)	N	(%)
30–39	2	(20.0)	46	(5.0)
40–49	11	(13.3)	57	(7.4)
50–59	14	(21.2)	52	(7.9)
60–69	11	(23.9)	163	(9.2)
Age-adjusted	40	(18.5)	326	(7.7)

Occupation-Specific Prevalence of Stones: Tennessee Workers

Occupation	Potential Uranium Exposure	Workers N	Prevalence Stones N	Prevalence Stones (%)
Uranium workers				
Production	(High)	95	19	(20)
Office	(Low)	20	5	(25)
Guards	(Low)	43	3	(7)
Dairy workers	(None)	50	13	(26)
Total		208	40	(19)

Discussion

Our principal finding is an increased risk of stones among males in eastern Tennessee, consistent with previous reports of a regional "stone belt."[1-3] At highest risk are persons with a positive family history whose urine is supersaturated with CAOX, a major ingredient of most stones.[4] Provided that our findings reflect regional occurrence, they suggest that stones are an important public health problem in the southeast, and provide some clues about etiology.

The lifetime risk of stones among the Tennessee workers is higher than in any other reported population. Risk by age 60 is 24 percent in our study, compared to 5 percent in Minnesota,[5] 7 percent in northern California,[4] and [illegible] percent in Uppsala, Sweden.[18] These regional differences are larger than can be explained by differences in the methods for estimating risk. The Swedish data, for example,[18] are based upon a prevalence survey of renal colic in 1781 men, age 60. The US NHANES survey,[13] using similar methods but inquiring about stones instead of renal colic, show a lifetime prevalence of 9.2 percent in White males ages 60–69. The lower risks observed in the California and Minnesota studies are based upon incidence rates of physician visits.[4,5] Incidence is converted to risk using the formula [$\text{risk} = 1 - e^{(-\text{Incidence} \times \text{Time})}$].

Several findings suggest that the risk of stones in our study is regional rather than occupational. First, a family history of stones affecting first degree relatives who have never worked at either plant was common in both stone-formers (52 percent) and those without stones (21.5 percent). This can be compared to Sweden, where 29.4 percent of stone-formers and 15.3 percent of nonstone-formers report a family history.[18] Second, the occurrence of stones in the uranium workers bears no relation to the intensity or duration of occupational exposure. Stones are slightly less common in workers with direct exposure to uranium and thorium than in unexposed workers. Third, stone occurrence is similar in the uranium and dairy workers. If stone occurrence were due to occupation it would be unlikely that two such different occupational settings would cause a nearly identical effect.

The geographic boundaries of the so-called "stone-belt" are poorly defined. Four southeastern states (North Carolina, South Carolina, Virginia, and Tennessee) rank high in terms of hospitalizations for stones as a proportion of all hospital admissions.[1] There are limitations to these hospital discharge data, however,[4,5] and no studies have collected population-based data on incidence in the southeast US. Such data would be extremely useful in mapping the occurrence of stones, defining high-risk areas, and identifying preventable risk factors.

No cause has yet been identified for increased stone-formation in the southeast. Various hypotheses include water softness,[19] sunlight and heat,[20] dietary consumption of meat,[16] and carbonated beverages.[21] All of these hypotheses are based upon ecological evidence and have not been tested in studies assessing both stone occurrence and the exposure of individuals to the proposed risk factors. Our study is the first to measure biochemical risk factors in urine in a high-risk population; as such it helps to narrow the range of hypotheses to be pursued in future research. The single most important factor contributing to CAOX supersaturation in our study was increased urinary calcium.

Twenty-five (20.7 percent) of participants in the medical study had clinical hypercalciuria (>300 mg per 24 hours).[16] We did not obtain a dietary history and cannot assess whether the increased uri-

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*Fitzsimmons SC: Self-reported kidney stone disease among Black and White adults: Prevalence and correlates in NHANES 11 (1976–1980). Presentation at 116th annual meeting of the American Public Health Association, Nov. 15, 1988, Boston.

TABLE 3—Demographic and Clinical Characteristics of Persons with and without Stones

Characteristic	Stones			No Stones			p value⁺
	Mean	(SD)	Range	Mean	(SD)	Range	
Age years*	53.5	(8.5)	38–71	51.7	(7.9)	38–72	—
Weight kg.**	92.8	(15.5)	69–135	85.4	(13.0)	56–120	(0.01)
% Born locally**	86.2			85.2			—
% Away ≥ 20 years**	9.7			19.4			(0.2)
% Ever use moonshine liquor**	25.9			61.5			(0.001)
% Family history of stones**	51.7			21.5			(0.001)
% Family history of infection**	20.7			5.7			(0.02)
% Personal history of infection**	41.1			21.5			(0.009)
% Smoking, current**	34.5			35.5			—
% Phenacetin usage**	0.0			7.2			—
Calcium, serum (mg/dl)***	9.8	(1.0)	9.3–10.6	9.6	(1.0)	9.2–10.6	(0.006)
Calcium urine (mg/dl)***	13.2	(2.2)	2.8–44.7	7.8	(1.9)	2.4–42.6	(0.001)
Calcium urine (mg/day)***	250	(1.5)	98–455	164	(2.0)	9–850	(0.0001)
Oxalate, urine (mg/day)***	47.6	(1.3)	29–74	41.2	(1.4)	13–83	(0.051)
Uric acid serum (mg/dl)***	6.3	(1.2)	4.2–8.9	6.5	(1.3)	3.5–9.8	—
Uric acid urine (mg/day)***	517	(1.8)	149–1434	504	(1.6)	125–1423	—
Urine volume (L/day)***	1.6	(1.4)	0.8–2.8	1.7	(1.6)	0.4–3.9	—
Calcium-oxalate index***	2.4	(1.7)	0.5–6.2	1.5	(1.8)	0.1–4.2	(0.0009)
Brushite index***	2.3	(2.4)	0.2–7.0	1.4	(2.7)	0.02–7.1	(0.048)

*Data on 208 persons completing core questionnaire.
**Data on 136 participants completing medical questionnaire.
***Geometric mean for 123 subjects providing blood and 24-hour urine.
⁺P-value stated only if p ≤ 0.20.

TABLE 4—Logistic Regression: Factors Predicting Kidney Stones*

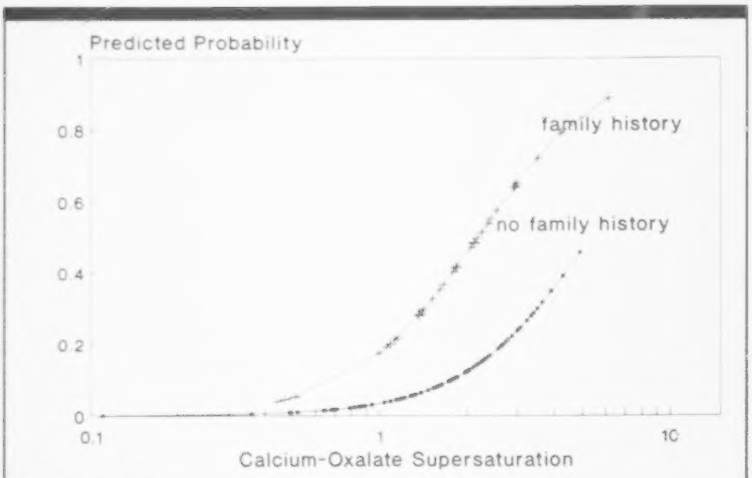
Risk Factors	Odds Ratio	SE	(95% CI)
Univariate Analyses			
Calcium-Oxalate Saturation Index			
Second tertile*	4.62	0.782	(1.0, 21.4)
Third tertile*	6.20	0.760	(1.4, 27.5)
Family history of stones	3.30	0.410	(1.5, 7.4)
Family history of infection	2.64	0.503	(1.0, 7.1)
Age 50–59 years	1.12	0.493	(0.4, 2.9)
Age ≥ 60 years	1.65	0.493	(0.6, 4.3)
Multivariate Analyses			
Calcium Oxalate Saturation Index			
Second tertile*	3.79	0.803	(0.8, 17.8)
Third tertile*	5.91	0.765	(1.3, 26.5)
Family History of Stones	3.51	0.425	(1.5, 8.1)
Age 50–59 years	1.62	0.505	(0.6, 4.4)
Age ≥ 60 years	2.21	0.505	(0.8, 5.9)

*Analysis based on 123 persons submitting blood and 24-hour urine sample.
**Mean scores for calcium-oxalate tertiles are 0.90, 1.79 and 3.01 for first, second, and third tertiles. Index is defined in text.

nary calcium was due to diet (milk, cheese, and meat) or to increased gastrointestinal absorption. However, we can exclude drinking water as a major source of calcium because municipal water supplies in the area have been monitored since 1982 and have unusually soft water due to low calcium.[22] Similarly, sunlight and its effect upon vitamin D cannot explain the measured values in our study, because the study was conducted in January. A second important determinant of CAOX was low urine volume. This cannot be attributed to incomplete urine collection, because it occurred in the presence of high 24-hour excretion of other stone risk factors. Low fluid intake is the most likely explanation for reduced urine volume and increased concentration. It is unclear why fluid intake should be low in these workers, although prohibition against alcoholic beverages in many areas of the southeast may play some role. Alcohol exerts a diuretic effect by suppressing the release of anti-diuretic hormone.[23] Consumption of moonshine liquor is negatively associated with stones in our study. The hypothesis that alcohol may protect against stones is plausible and deserves further evaluation.

Mildly elevated urinary oxalate (55–100 mg/24 hours)[24] occurred in 20 (16.5 percent) of the workers in our study. These levels are below those found in primary hyperoxaluria,[24] but may be consistent with dietary exposure. Foods commonly eaten in the southeast include several sources of oxalate (collards, mustard greens, kale, and ice tea).[25] Oxalate excretion contributed relatively little to urinary saturation with CAOX compared to hypercalcuria and low urine volume. However, the presence of oxalates in typical southern cuisine provides an intriguing hypothesis to be tested in future studies.

We cannot explain the low frequency of stones among the guards in our study. Guards were culturally different from the other workers. Twenty guards (46.5 percent) vs 10 (7.9 percent) of the other workers were retired military servicemen who had spent 20 or more years away from the area ($p < .0001$). Too few guards participated in the medical study to allow us to examine whether differences existed in their urine composition. The low participation occurred largely because guards



* defined in text

FIGURE 1—Probability of Stones by Family History and Calcium-Oxalate Supersaturation Index*

could not be released from duty to participate during working hours.

A limitation of our study is that the assessment of risk factors for stone-formation in urine is based upon a single 24-hour sample, obtained some years after the actual passing of stones. It is likely that excretion of most of these substances varies over time. Nevertheless, the strong association between current CAOX excretion and lifetime stone occurrence is consistent with the known importance of calcium and oxalate in stone formation. Excretion appears to be sufficiently stable to reflect a gradient of risk as current concentration increases. In summary, the study found an increased risk of urolithiasis and risk factors related to stone formation in male workers in eastern Tennessee. Future research should define the geographic range of increased kidney stone formation in the southeast and identify its causes. ☐

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Benzodiazepine Anti-anxiety Agents: Prevalence and Correlates of Use in a Southern Community

Marvin Swartz, MD, Richard Landerman, PhD, Linda K. George, PhD, Mary Lou Melville, MD, Dan Blazer, MD, PhD, and Karen Smith, PhD

ABSTRACT

Background: Benzodiazepine anti-anxiety agents are the most widely prescibed psychotherapeutic drugs in the United States today. Recent evidence, however, suggests that their use may be decreasing.

Methods: We examine the population prevalence and correlates of use of benzodiazepine anxiolytics at the Duke site of the NIMH-sponsored Epidemiologic Catchment Area project.

Results: Bivariate analysis of use patterns for the drugs revealed demographic predictors similar to those reported in previous studies: increased likelihood of use by the elderly, Whites, women, the less educated, and the separated or divorced. Use is also associated with symptoms of psychic distress, negative life events, use of health care services, and diagnoses of affective disorder, agoraphobia with panic, and panic disorder. Age, sex, race, education, and marital status remain associated with non-hypnotic benzodiazepine use in a logistic regression analysis.

Conclusions: Multivariate analyses of these data indicate that when potential confounding factors are controlled, age, sex, race, education, and marital status are significantly related to benzodiazepine anxiolytic use but the effects of sex and education are mediated by intervening variables. Implications of these findings are discussed particularly in relation to high levels of use in the elderly. (*Am J Public Health* 1991;81:592-596)

Introduction

Benzodiazepine anti-anxiety agents are the most widely prescribed psychotherapeutic drugs in the United States today.[1] First introduced in 1960, these drugs rapidly achieved a lead position in the prescription drug market,[2] stimulating public and professional debate over appropriate psychotropic drug use.[3] Recent evidence, however, suggests that the prevalence and patterns of psychotropic use, especially those of benzodiazepine anxiolytics, may be changing and resulting in decreased use.[4,5] The first detailed population survey of psychotropic drug use, the National Household Sample in 1970–71,[6-9] found that 22 percent of American adults had used prescription psychotropic medication during the year 1969–70, with higher use among women and the elderly. The most commonly used drug group, "minor tranquilizer/daytime sedatives," was used by 8 percent of men and 20 percent of women. Use was greater in Whites and in divorced, separated, and widowed persons and the elderly.[6] In 1979, the National Survey of Psychotherapeutic Drug Use documented, using weighted data, that an estimated total of 8 percent of men and 13.7 percent of women had used benzodiazepine anxiolytics in the previous year with the greatest use among respondents ages 50–64.[2,10] Recent surveys of the St. Louis area population[11] and Great Britain[12] found lower overall rates of use but similar sex and age usage profiles.

Previous studies have relied primarily on bivariate analyses when describing patterns of drug use, failing to take into account the intercorrelations among predictors, including demographic, diagnostic, symptoms, and service use variables. In the present study we examine the prevalence and patterns of benzodiazepine antianxiolytic drug use in the Piedmont region of North Carolina during 1982–83, utilizing logistic regression analysis, which allows prediction of benzodiazepine use while introducing controls for potential confounding and mediating variables.

Method

The present paper reports results from Wave 1 of the Piedmont Health Survey, one site of the five-site National Institute of Mental Health Epidemiologic Catchment Area program (NIMH-ECA).[13] The sampling frame for the Piedmont Health Survey (PHS) was a five-county area in north central North Carolina, consisting of one urban county and four contiguous rural counties. The survey randomly sampled all housing units from segments throughout the catchment area, using the Kish method,[14] yielding 3,798 interviews usable for this study; further details of the study have been published elsewhere.[15] In this paper analyses are based on weighted data, taking into account household probability selection, nonresponse, and the 1980 census demographic profile of adults in the five-county catchment area. For significance tests, the data were downweighted to the original number of subjects, but weighting adjustments remain.

Diagnoses were obtained through use of the Diagnostic Interview Schedule

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From the Psychiatric Epidemiology and Health Services Research Program, Department of Psychiatry, Duke University Medical Center. Address reprint requests to Marvin Swartz, MD, Box 3173, Duke University Medical Center, Durham, NC 27710. This paper, submitted to the Journal March 12, 1990, was revised and accepted for publication December 12, 1990.

(DIS), a highly structured interview designed for use by lay interviewers in epidemiological studies and capable of generating computer-based diagnoses for certain DSM-III disorders.[16] The diagnoses generated by the DIS were too numerous to include in a single regression equation. Preliminary analyses ascertained that an affective disorders summary measure (consisting of major depressive disorders, bipolar disorders, dysthymia, and bereavement), panic disorder, and agoraphobia with panic were most strongly related to benzodiazepine use. These measures were retained for subsequent analyses. The version of the DIS used in Wave 1 of this survey and other ECA sites did not include one anxiety disorders diagnosis, generalized anxiety disorder, employed at certain ECA sites in follow-up interviews.

In the present study psychotropic drug use refers to any use within the year prior to interview, regardless of amount or duration of use and was elicited by specific drug use probes aided by photographic drug identification cards. Other measures used for this analysis include a 29-item psychic distress scale analogous to the SCL-90,[17] and the number of negative life events during the year prior to the interview, measured by a scale of 20 major events of the type proposed by Holmes and Rahe.[18] Analyses of life events in the survey have been published elsewhere.[19,20] Interviewers also recorded the number of outpatient visits to physical and mental health providers during the six months prior to interview. The measure of outpatient health service utilization employed in this study is dichotomous, because the use vs non-use measure has more explanatory power than a continuous (number of visits) measure. Demographic variables used in this report include sex, age, race, education level, marital status, and urban vs rural residence. Income is excluded because preliminary analyses indicated that the relationship between anxiolytic use and income is weak.

The Appendix presents a six-stage causal model predicting the dichotomous measure of use vs nonuse of benzodiazepine anti-anxiety agents. A unidirectional causal process is assumed with variables in prior stages affecting those in subsequent stages. No causal direction is assumed for variables in the same stage. In the multivariate logistic regression analysis below, effects of the stage I variables are estimated first, and the equation is then expanded as groups of variables are entered by stage. This format yields an estimate of the effect of each variable net of all adjacent and prior variables and indicates whether the effects of prior variables are mediated by subsequent variables. The logistic regression procedure calculates maximum likelihood estimates for the parameters of a model that express the log odds of an event (here, drug use) as a simple linear model.[21,22]

Results

Table 1 presents the demographic profile for subjects participating in the survey.Table 2 focuses specifically on use of benzodiazepine anti-anxiety agents—predominantly diazepam and chlordiazepoxide—regardless of source. All other anxiolytics as well as all sedatives and hypnotics (benzodiazepine and other) are excluded. Table 2 presents a cross-tabular analysis of the weighted prevalence of benzodiazepine anxiolytic use among various demographic and other groups of interest. For two continuous variables, correlation coefficients are used: the correlation coefficient of drug use with number of negative life events is .09 ($p < .01$) and that for drug use with psychic distress symptoms is .20 ($p < .01$). Benzodiazepine anxiolytic use is more prevalent among women, older persons, Whites, and those with less education. The never married and married have significantly lower prevalence of use than the divorced and widowed. Respondents who have recently used either physical or mental health care services are much more likely to use these agents, as are those with a DIS diagnosis of affective disorder, panic disorder, or agoraphobia with panic.

Table 3 presents results of the logistic regression analyses. The "OR" coefficient in the table is the antilogged regression coefficient and estimates how changes in the independent variables multiply the odds of using benzodiazepine anxiolytics, holding other variables in the equation constant. Ninety-five percent confidence intervals are also given. The overall fit of the model is assessed by R statistic which is roughly analogous to the multiple correlation coefficient in ordinary least squares regression after correcting for the number of parameters estimated.[22]

TABLE 1—Demographic Profile of Community Subjects Participating in the Piedmont Health Survey (N = 3,798)

	No.[a]	%[b]	WTD[c]
Sex			
Male	1489	39	46
Female	2309	61	54
Race			
White	2399	63	62
Other[d]	1395	37	38
Age (years)			
18–34	1096	29	43
35–54	863	23	29
55–74	1414	37	23
75+	425	11	5
Marital Status			
Married with spouse	953	51	60
Separated/ divorced	480	13	11
Widowed	814	21	9
Never married	551	15	21
Educational Level (years)			
0–8	973	26	16
9–11	818	22	22
12–13	1062	28	34
14+	935	25	28
Residence			
Rural	1936	51	44
Urban	1962	49	56

[a] The total of some cells do not equal 3,798 because of missing data.

[b] The percents may not sum to 100 due to rounding.

[c] The weighted percentage adjusts the actual response percentage to take into account household probability selection, response rate, the elderly oversample and demographic characteristics of the 196,790 persons in the five-county area who were 18+ years of age. (Total five-county population, based on 1980 census data, was 269,863.)

[d] Other races are almost entirely Black.

In stage I, older persons and women are both more likely than others to use benzodiazepines, while Blacks are less likely. The gender effect decreases substantially in stages V and VI indicating that much of the initial sex difference is due to higher levels of psychic distress and health service utilization among women. The race effect changes very little across stages indicating no substantial effect of potential mediating variables. While the coefficients for age fluctuate across stages, they remain largely unmediated and substantial at stage VI. In stage II, increases across each of four levels of education (as indicated in Table 2) reduce the odds of benzodiazepine use by .9. This initial effect decreases in stages IV and V, indicating that those with more education use benzodiazepines less, in part, because they have fewer DIS diagnoses and psychic distress symptoms. Compared to the never married, the separated, divorced, and widowed are more likely to use benzodiazepine anxiolytics. These marital status effects are not mediated by varia-

TABLE 2—Bivariate Correlates of Benzodiazepine Anti-anxiety Agent Use, Piedmont Health Survey 1982–83

Variables	Total	Number Using Drug	WTD% Using Drug
Sex			
Males	1489	104	6.9
Females	2309	310	10.6[a]
Age (years)			
18–34	1096	50	3.8
35–54	863	99	11.5
55–74	1414	208	14.0
75+	425	57	13.6[a]
Race			
White	2399	317	11.5[a]
Non-White	1395	97	4.7
Marital Status			
Married	1953	197	9.3
Separated/divorced	480	70	11.4
Widowed	814	122	14.7
Never married	551	25	3.6[a]
Educational Level (years)			
0–8	973	124	11.3
9–11	818	111	11.8
12–13	1062	102	7.4
14+	935	77	7.2[a]
Residence			
Rural	1936	192	8.2
Urban	1862	222	9.4
DIS Diagnosis			
No Affective Dx	3638	364	8.2
1 Affective Dx	136	39	21.7
2+ Affective Dx	24	11	41.7[a]
No Agor. w/Panic Dx	3658	386	8.5
Agor. w/Panic Dx	40	10	33.1[a]
No Panic Disorder	3762	405	6.6
Panic Disorder Dx	22	6	39.2[a]
Health Service Utilization			
No Outpatient Physical Health Care (6 mos)	1619	84	3.9
Outpatient Physical Health Care (6 mos)	2150	325	13.2[a]
No Outpatient Mental Health Care (6 mos)	3450	301	7.0
Outpatient Mental Health Care (6 mos)	283	101	31.0[a]

[a] Significant at the .05 level or better. Where a particular set has more than two categories, significance test applies to the set.

bles entered in subsequent stages. Urban residence is not strongly associated with drug use. In stage III the number of negative life events is positively and substantially associated with drug use. Decreases in this effect in stages IV through VI indicate that the effect of negative life events is due primarily to the effects of life stress on DIS diagnoses, psychic distress symptoms, and increased health care utilization. In stages IV through VI, affective disorder, agoraphobia with panic, and panic disorder are all associated with increased benzodiazepine use, as are psychic distress and health care utilization. However as can be seen in the final model, effects of DIS diagnoses are mediated by psychic distress.

Discussion

Multivariate analysis of survey data indicates that when potential confounding factors are controlled, age, sex, race, education, and marital status are related to benzodiazepine anxiolytic use. Age, race, and marital status remain substantially unmediated by other factors while sex and education, previously thought to be predictors of anxiolytic use, exert their effects indirectly via their relationships with other predictors in subsequent model stages. The effect of education is mediated by intervening factors such as high levels of psychic distress or stressful life events. Women indeed use more anxiolytics, but probably because they either have or report more distress than men and/or are more likely to seek help. That is, net of the effect of increased levels of psychic distress and help-seeking among women, women are no more likely than men to receive and/or use anxiolytics, indicating less evidence for sex bias in prescribing practices and use than suggested in previous studies and the press.[11,23–25]

One of the most important findings in the present study relates to increased anxiolytic use in the elderly, many of whom are already medicated and/or cognitively impaired.[2] Benzodiazepine use in the elderly is a particular concern because the use of long half-life benzodiazepines such as the ones included in the present study is associated with an increased risk of hip fracture,[26] creating a potentially fatal complication of anxiolytic use in this population. High levels of use in the elderly raises concerns about the judicious use of these agents.

The differential rates of anxiolytic use by race, lower among Blacks even when predictors such as use of health services are controlled, cannot be explained further in the present analysis. These results do suggest possible racial bias in prescribing practices and/or different modes of help-seeking and alleviation of psychic distress among Whites vs Blacks. Prior studies have demonstrated that separated or divorced respondents have higher levels of life stress and psychic distress[9] and are more likely to use anxiolytics, presumably as a result of these stressors.[6,9] Our analyses suggest, however, that separated and divorced persons have a greater likelihood of drug use even net of negative life events and psychic distress, suggesting other explanatory or predispositional factors, not examined in the present analysis.

Finally these analyses raise questions about associations between symptoms, diagnosis, and appropriate psychotherapeutic drug use. There is evidence that the American public has become more conservative in its attitudes toward psychotropic drug use.[2,27] In fact, our preliminary data indicate that the majority of persons with high levels of anxiety in the community are not receiving psychotherapeutic medications; further analyses need to ex-

TABLE 3—Logistic Regression Effects Corresponding to Model in Appendix

	I OR (95% CI)	II OR (95% CI)	III OR (95% CI)	IV OR (95% CI)	V OR (95% CI)	VI OR (95% CI)
Intercept	0.04 —	0.05 —	0.03	0.03 —	0.02 —	0.01 —
Age 35–54	3.2 (3.8, 2.7)	2.7 (3.2, 2.2)	2.8 (3.4, 2.4)	3.0 (3.6, 2.5)	3.1 (3.8, 2.6)	3.1 (3.7, 2.5)
Age 55–74	4.0 (4.8, 3.4)	3.1 (3.7, 2.5)	3.4 (4.1, 2.8)	3.7 (4.5, 3.0)	4.0 (4.9, 3.3)	3.4 (4.2, 2.8)
Age 75+	3.6 (4.7, 2.8)	2.6 (3.5, 1.9)	2.9 (3.9, 2.1)	3.1 (4.3, 2.3)	3.4 (4.7, 2.5)	2.7 (3.7, 2.0)
Black	0.4 (0.4, 0.3)	0.4 (0.4, 0.3)	0.3 (0.4, 0.3)	0.3 (0.4, 0.3)	0.3 (0.4, 0.3)	0.3 (0.4, 0.3)
Female	1.5 (1.7, 1.3)	1.4 (1.6, 1.2)	1.4 (1.6, 1.3)	1.4 (1.6, 1.2)	1.3 (1.4, 1.1)	1.2 (1.4, 1.0)
Education		0.9 (0.9, 0.8)	0.9 (0.9, 0.8)	0.9 (1.0, 0.8)	1.0 (1.0, 0.9)	0.9 (1.0, 0.9)
Married		1.4 (1.8, 1.1)	1.5 (1.9, 1.1)	1.5 (1.8, 1.2)	1.5 (1.9, 1.2)	1.6 (2.0, 1.2)
Separated/divorced		2.2 (2.9, 1.6)	2.1 (2.8, 1.6)	2.1 (2.8, 1.6)	2.1 (2.8, 1.6)	2.3 (3.0, 1.7)
Widow		1.8 (2.4, 1.3)	1.7 (2.4, 1.3)	1.8 (2.4, 1.3)	1.7 (2.4, 1.3)	2.0 (2.7, 1.5)
Urban		1.3 (1.4, 1.1)	1.2 (1.4, 1.1)	1.2 (1.3, 1.0)	1.1 (1.2, 1.0)	1.1 (1.3, 1.0)
#Life stress			1.5 (1.6, 1.4)	1.4 (1.5, 1.3)	1.2 (1.3, 1.1)	1.1 (1.2, 1.0)
#Affect Dx				2.2 (2.6, 1.8)	1.2 (1.5, 1.0)	1.1 (1.3, 0.9)
Dx-agor/pan				2.8 (4.3, 1.8)	1.3 (2.0, 0.8)	1.3 (2.1, 0.8)
Dx-panic				4.6 (7.9, 2.6)	2.3 (4.0, 1.3)	1.3 (2.4, 0.8)
Psyc distress					1.2 (1.3, 1.2)	1.2 (1.2, 1.2)
Outpat phys						2.7 (3.2, 2.3)
Outpat ment						2.3 (2.7, 1.9)
Model:						
R	.26	.27	.30	.32	.36	.41

plore whether they are receiving other psychotherapeutic interventions. There is also a strong association between a DIS diagnosis of affective disorder and use of benzodiazepine anxiolytics. Concurrent use of antidepressants is not examined in the present study. However, Craig, *et al*, also found that surveyed women taking minor tranquilizers and sedatives included significantly more respondents with high depression scores, and that depressed respondents of both sexes tended to be taking anxiolytics or sedatives, and not antidepressants.[28] Beardsley, *et al*,[29] found that primary care practitioners provided 72 percent of all anxiolytic drug visits, but documented a psychiatric diagnosis in a minority of anxiolytic drug visits, raising questions about the specificity of anxiolytic use in primary care settings. Although anxious and depressive symptoms are often admixed, the present findings of high anxiolytic use among depressed respondents suggest that depressive symptoms are underrecognized. Major depressive episodes usually require specific antidepressant medication and psychotherapy, and treatment with benzodiazepine anxiolytics in the absence of more specific antidepressant therapy may postpone therapeutic response or even exacerbate symptoms.[23,30,31]

We have examined the current prevalence of benzodiazepine anxiolytic use in a Southern community, demonstrating that age, sex, race, education, and marital status are significantly associated with drug use, but that the effects of sex and education are mediated by intervening variables. In addition we have found benzodiazepine anxiolytic use to be significantly associated with several DIS diagnoses although the effects of diagnoses are largely mediated by psychic distress and use of health services. Future research on this population will focus on duration and frequency of anxiolytic use, use of other psychotropic drugs, and use of potentially interacting drug combinations. Finally, more detailed DIS diagnostic information gathered at follow-up of the survey cohort, including generalized anxiety disorder, will allow closer examination of the associations between benzodiazepine anxiolytic use and DIS diagnoses of affective and anxiety disorders. ☐

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APPENDIX

A Six-Stage Causal Model Predicting Benzodiazepine Use

I	II	III	IV	V	VI
Age	Education	Life Events	Affective Disorders	Psychic Distress	Utilize Outpatient Physical Health Services
Race	Marital Status		Agoraphobia with panic		
Sex	Urban Residence		Panic Disorder		Utilize Outpatient Mental Health Services
			Other Psychiatric Diagnoses		

The Influence of Three Mass Media Campaigns on Variables Related to Adolescent Cigarette Smoking: Results of a Field Experiment

Karl E. Bauman, PhD, John LaPrelle, PhD, Jane D. Brown, PhD, Gary G. Koch, PhD, and Connie A. Padgett, BA

ABSTRACT

Background: This paper reports findings from a field experiment that evaluated mass media campaigns designed to prevent cigarette smoking by adolescents.

Methods: The campaigns featured radio and television messages on expected consequences of smoking and a component to stimulate personal encouragement of peers not to smoke. Six Standard Metropolitan Statistical Areas in the Southeast United States received campaigns and four served as controls. Adolescents and mothers provided pretest and posttest data in their homes.

Results and Conclusions: The radio campaign had a modest influence on the expected consequences of smoking and friend approval of smoking, the more expensive campaigns involving television were not more effective than those with radio alone, the peer-involvement component was not effective, and any potential smoking effects could not be detected. (*Am J Public Health* 1991; 81:597-604)

Introduction

There is general agreement that cigarette smoking is the major preventable cause of death in the United States and that efforts to prevent smoking onset should be directed toward young people.[1]

Few studies have evaluated the influence of broadcast media on adolescent smoking.[2] Lewitt, Coate, and Grossman concluded that youth smoking rates were reduced by messages broadcast on radio and television.[3] The awareness of negative personal and social consequences of smoking increased among Minnesota adolescents after the introduction of campaigns that involved television, radio, and billboards.[4] Worden, *et al.*, found fewer reports of friend smoking, but no less smoking, among adolescents in a rural Vermont county that received a television campaign than in four matched control counties.* Flay, *et al.*, found no effects on smoking intention or behavior by a coordinated television and school-based prevention program.[5] Research on mass media campaign effects for behaviors other than smoking rarely involves adolescents (see, for example, Rice and Atkin[6]).

The purpose of this research was to evaluate the impact of mass media campaigns intended to prevent adolescent cigarette smoking.

Methods

The Mass Media Campaigns

The three mass media campaigns evaluated in this research are described in detail elsewhere.[7,8] They were developed with the guidance of behavioral science theory and research, one year of extensive formative research, and the principle that they could be implemented readily throughout the United States.

One campaign used eight 30-second radio messages that focused on seven expected consequences of smoking that are related to whether young people become regular smokers. We identify this campaign, and the areas in which it was implemented, as RADIO.

The second campaign (RPEER) was similar but also included a 60-second radio message that invited persons 12 to 15 years old to enter the "I Won't Smoke Sweepstakes." The purpose of the sweepstakes was to obtain names and addresses. We mailed entrants a brochure that asked them to talk to their friends about not smoking, to encourage their friends to pledge not to smoke, and to have their friends enter the sweepstakes. We offered a $20 incentive for recruiting five or more

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*Worden JK, Flynn BS, McAuliffe TL, Sweeney RR, Secker-Walker RH: Using television to reach rural children with non-smoking messages. Paper presented at the American Rural Health Association Conference, Jeffersonville, VT, June 1982.

From the University of North Carolina, Chapel Hill, where Dr. Bauman and Ms. Padgett are with the School of Public Health, Department of Health Behavior and Health Education; Dr. LaPrelle is with the Department of Psychology; Dr. Brown is with the School of Journalism; and Dr. Koch is with the School of Public Health, Department of Biostatistics. Address reprint requests to Karl E. Bauman, PhD, Department of Health Behavior and Health Education, School of Public Health, the University of North Carolina, Chapel Hill, NC 27599-7400.
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entrants. The brochure also was mailed to friends they recruited.

The third campaign (RTVPEER) was similar to the second but included television broadcast of the sweepstakes offer and only three of the expected consequences messages.

The expected consequences messages were broadcast throughout November 1985, January 1986, and April 1986. The sweepstakes offer was broadcast during November 1985 and the brochures to encourage personal peer contact were mailed from January 1986 through February 1987. To simulate implementation of a national campaign, an advertising agency purchased broadcast time. The agency placed the messages at times expected to reach 75 percent of the intended audience four times during each of three four-week periods.

Design

The study design involved ten Standard Metropolitan Statistical Areas (SMSAs). Two SMSAs received each treatment condition and four served as CONTROL.

To minimize heterogeneity among study areas, only the 81 SMSAs in the Southeast United States were eligible for study. Homogeneity was increased further by excluding SMSAs with populations of fewer than 200,000 (n = 22) and more than 500,000 (n = 21) inhabitants, median age above 36 (n = 4), more than 90 percent White (n = 12), more than 22 percent with four or more years of college and ages 25 or older (n = 4), and density less than 100 or more than 300 per square mile (n = 2). We randomly eliminated SMSAs with overlapping broadcast areas until 10 SMSAs, located evenly throughout the Southeast United States, remained. We divided the Southeast into north, middle, and south, and randomly allocated the SMSAs to treatment conditions within each area. Cost and legal restrictions required two reassignments within area after random allocation. The final distribution of study SMSAs by treatment condition is CONTROL (Chattanooga, Tennessee; Columbia, South Carolina; Jackson, Mississippi; Savannah, Georgia), RADIO (Lakeland, Florida; Macon, Georgia), RPEER (Montgomery, Alabama; Roanoke, Virginia), and RTVPEER (Lexington, Kentucky; Mobile, Alabama).

Data Collection

Cluster sampling procedures identified probability samples of households within each SMSA in 1985.** Households were screened for adolescents ages 12 to 14. For the pretest surveys, conducted from April 1, 1985 through October 13, 1985, interviewers attempted to obtain data from all eligible adolescents and their mothers in the households. When more than one adolescent age 12 to 14 resided in a household, one was randomly selected to serve as a subject. Of the 2,534 adolescent subjects eligible for study in all SMSAs, 2,102 (83.0 percent) participated at pretest. The same subjects were asked to provide data again from April 16, 1987 to October 29, 1987, 11 to 17 months after the broadcasts ended and two to eight months after the mailings for the peer-involvement component ended. Of those who participated at pretest, 1,637 also participated at posttest and are included in the analyses reported here. Some pretest subjects could not be located at follow-up or had substantial missing information and are therefore lost to study. The numbers of subjects across SMSAs range from 132 to 232.

To assess overall attrition bias, we compared subjects who provided data only at pretest with those who also provided data at posttest on the intervening and smoking variables measured at pretest. The two groups did not differ on variables measuring smoking experimentation, subjective expected utility, smoking intention, friend approval of smoking, or friend encouragement of not smoking. Subjects lost to study, compared to subjects who provided both pretest and posttest data, were significantly more likely to be recent smokers (5.0 versus 3.1 percent), to be regular smokers (5.6 versus 3.1 percent), and to have high mean smoking intensity (1.22 versus 1.16).

While collecting posttest data from the panel subjects, we also gathered data from a cross-sectional sample of adolescents 14 to 16 years old and their mothers. Procedures were similar to those used to select subjects at baseline for the panel sample but we used clusters other than those in the panel and fewer subjects. We refer to this as the cross-sectional sample to distinguish it from the panel sample. The cross-sectional sample increased sample size and allowed assessment of testing effects.[9] Of the estimated 1,759 subjects eligible for the cross-sectional sample, 1,216 provided data.

Adolescent subjects averaged one hour to complete a self-administered questionnaire in their homes and provided alveolar breath and saliva samples to measure tobacco use. The mothers of most adolescents also completed a 25-minute self-administered questionnaire and provided alveolar breath samples.

Measures

The variables are listed in the Appendix. Data from the mother questionnaire measured parent education. All other measures are from data provided by the adolescents. Their derivation follows.

Adolescent subjects indicated the likelihood they would experience each of the seven expected consequences of smoking featured in the campaigns (bad breath, difficulty concentrating, loss of friends, trouble with adults, loss of appetite, increased fun, and increased relaxation) and how much they would like or dislike each consequence. We multiplied the responses for each consequence and summed the products to indicate the degree to which more positive or negative consequences of smoking were expected, with the more positive score favoring smoking. This measures subjective expected utility (SEU) for smoking (smoking SEU).[10] SEU for not smoking (nonsmoking SEU) was measured in the same way but the consequences were for not smoking rather than for smoking; more positive scores favored not smoking. A combined SEU (total SEU) was derived by subtracting nonsmoking SEU from smoking SEU; more positive scores on total SEU indicate greater preferences for smoking.

Decreased friend approval of smoking and increased friend encouragement of not smoking were hypothesized to be direct consequences of the peer-involvement component. To measure friend approval, subjects were asked how many of their three best friends would approve of their smoking. To measure friend encouragement, subjects were asked how often someone their own age had encouraged them to be a nonsmoker; response categories ranged from never to 10 or more times. To measure smoking intention, subjects were asked how sure they thought they would smoke in the next two years, with response categories ranging from certain to smoke to certain not to smoke.

Subjects provided alveolar breath samples to measure carbon monoxide (CO); a value ≥ 9.0 ppm indicated recent smoking.[11] Subjects were asked whether they had ever puffed on a cigarette. Sub-

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**Research Triangle Institute: Survey of Young Adolescent Smoking Behavior: Final report, 1986.

jects also were asked about their smoking intensity using a single questionnaire item that measured recency, frequency, and amount smoked.[12,13] The values ranged from never smoked a cigarette to currently smoke more than two packs a day. We also used the intensity measure to classify subjects as regular smokers if they usually smoked weekly or more often. To increase the validity of self-reports of smoking, subjects were told before beginning the questionnaires that the alveolar breath and saliva samples they gave would be used to measure their smoking.[14,15]

Analysis

Baseline data were used to create subsamples of subjects who were not regular smokers. Of the 1,637 panel subjects, the nine who did not report whether they had puffed on a cigarette were excluded from all analyses. The remaining subjects were divided into those who had not (N = 973) and those who had (N = 655) puffed on a cigarette. Of the adolescents who had not puffed on a cigarette, 22 with missing information on CO or with CO ≥ 9 ppm were excluded; the remaining 951 subjects are nonexperimenters. Of the adolescents who had puffed on a cigarette, 127 who were regular smokers, had CO ≥ 9 ppm, or had missing information were excluded; the remaining 528 are experimenters. The total sample includes experimenters (N = 528), nonexperimenters (N = 951), regular smokers (N = 48), and subjects with insufficient information for classification (N = 110).

Each SMSA was assigned average scores based on the information provided by subjects within the SMSA. Multiple regression analysis for repeated measures was the principal statistical method for inferences concerning treatment differences for changes over time by SMSA, i.e. treatment by time interaction.[16] We did not include control variables in analyses using the SMSA as the unit because these units had small sample size (n = 10), were randomized, and were relatively homogeneous with respect to previously stated population characteristics. Analyses with the individual as the unit used logistic regression for dichotomous dependent variables and linear regression for continuous dependent variables. These latter analyses examined effects for treatment, SMSA nested in treatment, and control variables from the pretest. We included the control variables related to an intervening or dependent variable and to pretest differences between treatments, attrition bias, or missing data substitution. For an intervening or dependent variable at posttest, we controlled for the corresponding pretest measure of the variable.

Results

Independent Variables

Message reach and frequency were estimated by the advertising agency using standard estimating procedures with Arbitron and Nielsen data for radio and television, respectively.[17] The expected consequences messages reached 81 percent of the intended audience an average of 4.5 times during each of the three four-week periods that they were broadcast. The sweepstakes offer in RPEER reached 79 percent of the audience an average of four times. In RTVPEER, messages reached 90 to 94 percent of the intended audience by a combination of radio and television, and they were seen or heard an average of 14 times during the entire period of broadcast. Young people ages 12 to 15 years surveyed by telephone the month after the final campaign messages were broadcast reported higher levels of reach.[18]

Based on sweepstakes entry and US Census data, 21 to 24 percent of the adolescents 12 to 15 years old in RTVPEER entered the sweepstakes; 97.8 percent of them responded to the encouragement of others rather than to the broadcast messages.[8] Nine to 15 percent of the adolescents in RPEER entered the sweepstakes, and 94.8 percent of them responded to their friends rather than to the broadcast.[8]

Intervening Variables

Between pretest and posttest, all intervening variables changed significantly in the expected direction of becoming more favorable toward smoking (Table 1). The p values in Table 1 for the time by treatment interactions indicate whether the change in the intervening variables between 1985 and 1987 varies significantly by treatment. There were significant time by treatment interactions for smoking SEU (nonexperimenters and total sample), nonsmoking SEU (experimenters only), and total SEU (nonexperimenters and total sample).

In Table 2, we show the 1985 and 1987 values of smoking SEU for the total sample, by treatment condition, and measures of treatment effect for change between 1985 and 1987. Smoking SEU increased less in RADIO and RTVPEER than in CONTROL, increased by the same amount in RADIO and RTVPEER, and increased by the same amount in RPEER and CONTROL. These findings suggest that RADIO and RTVPEER had the intended effects of decreasing smoking SEU. These effects of generally 4.0 correspond to about one-third of the baseline standard deviation of smoking SEU, which is comparable to a shift between the median and the 60th percentile. Thus, the campaign effect size is modest rather than large. The findings were similar for the other SEU measures and samples for the statistically significant treatment by time interactions shown in Table 1.

The findings in Table 2 suggest that the peer-involvement component in RPEER might have had undesirable effects, but this seems unlikely given the lower participation in the peer-involvement component in response to RPEER than to RTVPEER. However, to address the possibility of undesirable effects attributable to the peer-involvement component, we disaggregated subjects within RPEER and RTVPEER according to their participation in the peer-involvement campaign. Participants had entered the sweepstakes or reported on their posttest questionnaires that they had been personally contacted to enter the sweepstakes; all other subjects are nonparticipants. We conducted analyses similar to those shown in Tables 1 and 2 using the intervening variables as dependent variables because the campaigns were designed to influence them directly. Within both RPEER and RTVPEER areas, participants did not differ from nonparticipants in their 1985 to 1987 change in SEU or friend approval of smoking, but participants were more likely than nonparticipants to report an increase in encouragement by friends not to smoke. These findings suggest that the peer-involvement component did not produce undesirable effects. The absence of other treatment conditions, such as one with both radio and television broadcast of expected consequence messages and no peer involvement component, precludes analyses that might clarify these findings.

Changes in the SMSAs between 1985 and 1987 other than the campaigns might have produced the higher increase in smoking SEU in RPEER than in RTVPEER and RADIO shown in Table 2. For example, if adolescent use of smokeless tobacco increased more in RPEER than in RADIO and RTVPEER and adolescent use of smokeless tobacco is related to smoking SEU, then the differential change in smokeless tobacco use might explain the smoking SEU patterns. Using the individual as the unit of analysis, we identified intervening

TABLE 1—Multiple Regression Analysis for Repeated Measures of Time, Treatment, and Time by Treatment Effects for Intervening Variables

		Means		p		
Variables	Subsample	1985	1987	Time	Treatment	Time by Treatment
Smoking SEU	Nonexperimenters	−20.41	−16.41	.001	.540	.014
	Experimenters	−15.62	−12.61	.037	.896	.391
	Total sample	−17.88	−14.72	.004	.748	.061
Nonsmoking SEU	Nonexperimenters	21.79	17.79	.001	.928	.386
	Experimenters	18.27	12.91	.001	.107	.029
	Total sample	19.96	15.65	.001	.760	.321
Total SEU	Nonexperimenters	−42.21	−34.46	.001	.772	.028
	Experimenters	−33.89	−25.47	.002	.373	.120
	Total sample	−37.84	−30.33	.001	.632	.092
Friend approval of smoking cigarettes	Nonexperimenters	.33	.65	.001	.700	.003
	Experimenters	.90	1.14	.010	.836	.257
	Total sample	.60	.89	.001	.860	.078
Friend encouragement of not smoking cigarettes	Nonexperimenters	1.06	1.28	.023	.196	.742
	Experimenters	1.21	1.48	.002	.341	.288
	Total sample	1.14	1.39	.004	.240	.778
Intention not to smoke cigarettes	Nonexperimenters	4.62	4.42	.001	.975	.304
	Experimenters	4.23	3.91	.005	.865	.449
	Total sample	4.40	4.16	.001	.896	.248

Notes: SEU = subjective expected utility
n = 10
Source: Panel sample

TABLE 2—Descriptive Statistics, Measures of Effect, and Treatment Comparisons for Smoking Subjective Expected Utility for Total Sample

						Treatment and Control Change	
Treatment	n	Statistics	1985 Mean	1987 Mean	1985 vs 1987 Change	Difference	(95% CI)
CONTROL	4	Estimate	−18.74	−14.16	4.58	Reference	
		SE	.70	.62	.93		
RADIO	2	Estimate	−16.64	−16.28	.36	−4.22	(−8.16, −.28)
		SE	.99	.87	1.31	1.61	
RPEER	2	Estimate	−19.22	−13.80	5.42	0.84	(−3.10, 4.78)
		SE	.99	.87	1.31	1.61	
RTVPEER	2	Estimate	−16.08	−15.20	.88	−3.70	(−7.64, .24)
		SE	.99	.87	1.31	1.61	
Comparison among Groups	p-value		.14	.24	.06	.06	

SE = Standard Error

variables other than SEU, and the control variables as measured at pretest, that were related to nonsmoking SEU, smoking SEU, or total SEU at posttest. They were dependent variables in analyses conducted to identify statistically significant time by treatment interactions that parallel the findings involving smoking SEU and total SEU. There were only two statistically significant and parallel interactions, but control for the involved variables did not account for the patterns in Table 2. The greater increase in SEU in RPEER than in RADIO and RTVPEER remains unexplained.

There were significant time by treatment interactions for friend approval of smoking for nonexperimenters and for the total sample (Table 1). The 1985 and 1987 values of friend approval for the total sample, by treatment condition, and measures of treatment effect for change between 1985 and 1987 are in Table 3. These findings suggest that RADIO caused less increase in friend approval of smoking relative to the control condition. The estimated effect of RADIO corresponds to a modest reduction of 7 percent in the number who reported that one or more of their friends approve of smoking. The findings were the same for nonexperimenters.

There were no significant time by treatment effects that involved smoking intention or friend encouragement of smoking.

Cigarette Smoking

Table 4 shows the findings with cigarette smoking as dependent variables. The means of all smoking variables changed significantly between 1985 and 1987 in the expected direction of more smoking. However, none of the p values for the time by treatment interactions were statistically significant. The data are consistent with the conclusion that the campaigns did not influence smoking.

There were substantial within-treatment differences in smoking between SMSAs that must be recognized because they precluded the ability to identify campaign effects for smoking if there were any to be

TABLE 3—Descriptive Statistics, Measures of Effect, and Treatment Comparisons for Friend Approval of Smoking Cigarettes for Total Sample

						Treatment and Control Change	
Treatment	n	Statistics	1985 Mean	1987 Mean	1985 vs 1987 Change	Difference	(95% CI)
CONTROL	4	Estimate	.56	.87	.31	Reference	
		SE	.05	.06	.05		
RADIO	2	Estimate	.72	.81	.09	−.22	(−.43, −.01)
		SE	.07	.09	.07	.09	
RPEER	2	Estimate	.58	.99	.41	.10	(−.11, .31)
		SE	.07	.09	.07	.09	
RTVPEER	2	Estimate	.60	.90	.30	−.01	(−.22, .20)
		SE	.07	.09	.07	.09	
Comparison among Groups	p-value		.39	.56	.08	.08	

SE = Standard Error

TABLE 4—Multiple Regression Analysis for Repeated Measures of Time, Treatment, and Time by Treatment Effects for Dependent Variables

		Means		p		
Variables	Subsample	1985	1987	Time	Treatment	Time by Treatment
Smoking experimentation	Nonexperimenters	—	.37	—	—	.906
	Total sample	.40	.57	.002	.410	.637
Regular smoking	Nonexperimenters	—	.05	—	—	.782
	Experimenters	—	.18	—	—	.907
	Total sample	.03	.14	.001	.916	.148
Recent smoking	Nonexperimenters	—	.03	—	—	.563
	Total sample	.03	.08	.001	.762	.157
Smoking intensity	Nonexperimenters	—	1.23	—	—	.910
	Total sample	1.16	1.50	.001	.876	.778

Note: n = 10
Source: Panel sample

detected. We began the study with the assumption that 4 percent of adolescent subjects in each study SMSA would be regular smokers and that SMSA changes in smoking within treatment condition would be similar. About 4 percent of our subjects were regular smokers in 1985. However, the range across SMSAs was .6 percent to 5.2 percent. The variation across SMSAs in the other smoking measures, and changes in smoking, also varied substantially across SMSAs. Most importantly, smoking varied substantially among SMSAs within treatment condition. Table 5 shows the treatment-control differences in smoking and the intervening variables that would have been required to achieve statistical significance at the .10 level for powers of .50 and .80 given the SMSA variation in smoking within treatment conditions in our data; the actual differences are also shown. The relatively large error variance for SMSAs within treatment groups clearly precluded detecting campaign effects for smoking. Therefore, whether the campaigns influenced smoking cannot be determined with acceptable certainty from our data.

Related Findings

We also used the 1987 cross-sectional data to examine campaign effects. For each SMSA, we assigned pretest values of the intervening and dependent variables in the panel data to the cross-sectional data. We then conducted multiple regression analysis for repeated measures with a sample size of 20 for the analysis: 10 panel and 10 cross-sectional SMSAs. The main statistic of interest was time by treatment interaction. The conclusions about campaign effects were generally the same as those when only the panel data were used.

Having subjects with and without baseline measures allowed us to identify effects due to the pretest. To assess this, as part of the analyses described in the preceding paragraph, we evaluated the time by treatment by data source (panel versus cross-sectional) interactions involving each intervening and dependent variable. The findings suggest that there were no testing effects.

We also conducted analyses to determine whether campaign effects for smoking varied across selected subgroups by testing for time by treatment by subgroup interactions with multiple regression analysis for repeated measures. The 10 variables used to create the subgroups, each divided at the median except when categorical, were: gender, age, smoking by best friend, smoking by mother, school curricula on smoking, total SEU, alcohol use, parent education, number of hours spent listening to the radio, and number of hours spent watching television. This analysis revealed no pattern that could be interpreted as campaign effects for smoking, and the conclusions for intervening variables remained the same.

TABLE 5—Required and Observed Differences between Treatment and Control

	Required[a] Power		Observed		
Variables	.50	.80	RADIO	RPEER	RTVPEER
Nonexperimenters					
Smoking experimentation	.19	.25	.01	.06	.00
Regular smoking	.06	.07	.02	.02	.00
Recent smoking	.03	.04	.01	.02	.02
Smoking intensity	.21	.28	.01	.03	.06
Total SEU[b]	4.38	5.79	8.96	.13	5.07
Smoking SEU[b]	2.63	3.48	6.34	.60	4.22
Nonsmoking SEU	2.99	3.96	2.63	.37	.86
Smoking intention	.19	.25	.04	.21	.01
Friend approval of smoking[b]	.08	.10	.27	.04	.00
Friend encouragement of not smoking	.38	.50	.12	.21	.01
Total Sample					
Total SEU[b]	5.25	6.94	5.86	2.38	5.29
Smoking SEU[b]	3.05	4.04	4.04	.98	3.60
Nonsmoking SEU	3.04	4.02	1.73	1.44	1.66
Smoking intention	.19	.25	.02	.20	.02
Friend approval of smoking[b]	.16	.22	.21	.09	.01
Friend encouragement of not smoking	.29	.38	.01	.15	.05

Note: SEU = subjective expected utility

a) Differences required to identify effects 50 and 80 percent of the time given study inter-SMSA variance.

b) Overall ANOVAs significant at alpha = .10 (n = 10).

Finally, we conducted analyses of the relationship between treatment and the intervening and dependent variables when using individuals rather than SMSAs as the unit of analysis. This procedure could artificially increase statistical detection of differences by ignoring intra-SMSA correlation among subjects.[19] In these analyses, we controlled for the pretest value of each control variable that was related to the posttest value of the intervening or dependent variable, the pretest value of the intervening or dependent variable being examined, and variables indicating attrition bias and effects from missing-data substitution. The findings were generally the same as when the SMSA was the unit of analysis. SEU effects were limited to RADIO and RTVPEER and no effects for cigarette smoking were detected.

Discussion

SEU and friend approval of smoking appear to have been influenced by the media campaigns. RADIO was as effective for these variables as any of the other more expensive campaigns. The peer-involvement component was not effective. We detected no effect for reducing the onset of smoking, and any potential effect for smoking remains to be demonstrated.

The substantial variation in smoking within treatment conditions is a major methodological limitation of this research because it precluded determining whether the campaigns influenced cigarette smoking. A consideration supporting the possibility of some campaign influence on smoking is that the campaigns appear to have influenced SEU and friend approval of smoking, variables that are related to the onset of smoking.[7,8] In addition, the campaigns featured characteristics that media scholars consider necessary for campaign effects, such as extensive formative research, content based on behavioral science theory, high quality message production, actors with characteristics admired by the audience, and repeated exposure to the campaigns by most of the intended population.[20-22] Earlier studies have found that mass media campaigns can influence behavior.[6]

Some contend that substantial impact by mass media campaigns requires their implementation with other types of interventions, such as community organization efforts or school-based curricula.[2,23] Our study was designed to determine whether the media by themselves are influential. This is the most common and practical application of mass media in public health and, unlike multiple-component approaches, is capable of ready distribution on a national level. However, this also precludes our knowing whether adding other types of prevention programs to the campaigns we designed and implemented would have produced larger effects.

To increase homogeneity of SMSAs within and across treatment conditions, we limited our study to SMSAs in the Southeast US, included only SMSAs similar on social and demographic characteristics, stratified the Southeast by area, and randomly allocated SMSAs to treatment conditions within area. Still, substantial variation in adolescent smoking precluded our determining whether there were campaign effects for smoking. We might have avoided the problem by studying a much larger sample of SMSAs and perhaps matching them for baseline smoking behavior before random allocation to treatment conditions.

Information on other interventions that might have been implemented to influence adolescent smoking in our study areas would have been useful for attempting to explain and control the SMSA variation in smoking. We assume that school-based programs would have been the only potentially significant organized activities implemented to reduce adolescent smoking during the time of our study; we controlled for school curricula and many other variables at baseline when using the individual as the unit of analysis. The substantial variation in adolescent smoking by SMSA, and our findings on effect size for intervening variables, suggest that more powerful interventions than the mass media campaigns we implemented might be necessary to demonstrate community-level behavioral effects for programs designed to prevent adolescent smoking.

We designed the peer-involvement component to influence perception of friend approval of smoking. However, the findings suggest that only RADIO, which did not have a peer involvement component, influenced friend approval. Perhaps RADIO influenced peer approval because two of the expected consequences messages focused on peers. Or, perhaps the radio messages were projected to their friends because they are like the young people who recorded them. The reason this effect did not also occur in response to RPEER and RTVPEER was not determined.

Given the centrality of personal influence to many behavioral science and mass communication theories, it is significant that the two campaigns with the peer-involvement component, RPEER and RTVPEER, were not more effective than RADIO. Perhaps the personal contact produced by the peer-involvement component was too superficial for impact, perhaps this approach is less powerful than commonly believed, or perhaps effects would have appeared for younger or older people. We do not have the data required to assess these possibilities.

Implementing the campaigns in 1985 for a broadcast reach of 75 percent of the United States population 12 to 17 years old under the conditions of our field experiment, excluding development and research expenses, would have cost $1,843,000 for RADIO, $12,866,000 for RPEER, and $28,944,000 for RTVPEER. Although radio is the less expensive medium, television is the broadcast medium that receives the most attention by those who design, implement, and evaluate mass media campaigns. Perhaps the glamour associated with television explains why there have been few recent studies of health-oriented radio campaigns. Radio use increases relative to television use during adolescence,[24] and our findings suggest that radio should be given more attention in campaigns to influence adolescent smoking and by research on campaign effectiveness.

Although findings that demonstrate program impact for intervening variables such as SEU and friend approval but not for behavior can sometimes be dismissed readily, one can also argue that they merit attention. SEU for smoking and friend approval for smoking have been considered by many to be precursors of smoking.[7,8] Our findings suggest that RADIO did influence SEU and friend approval. These findings, together with the capability of radio campaigns to be implemented readily on a national scale or to other large populations at lower cost than virtually any other type of intervention, support its potential role for health promotion among adolescents. Whether adolescent smoking would be influenced by our radio campaign or by others, however, has yet to be shown. The failure of the more expensive components, television broadcast and peer involvement, to influence the intervening variables suggests that these components might be unworthy of further consideration, at least as they were implemented in our research.

Our findings cannot be generalized to all applications of the mass media to smoking prevention. We might have found other effects if we had done things differently, such as focused on younger or older adolescents, implemented the campaigns in smaller or larger communities, or also implemented smoking prevention curricula in schools. The effects also might have been different if we had used other theories for guidance and if the style and format of our messages had been different. Flay, *et al*,[25] and Worden *et al*,[26] describe other approaches to using the mass media to prevent adolescent smoking. Finally, our findings cannot be generalized to the many public service campaigns that rely solely on donated broadcast time and fail to reach most of their intended audiences. ☐

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APPENDIX—Intervening, Dependent and Control Variables

Intervening

- Smoking subjective expected utility
- Nonsmoking subjective expected utility
- Total subjective expected utility
- Friend approval of smoking
- Friend encouragement of not smoking
- Smoking intention

Dependent

- Smoking experimentation
- Smoking intensity
- Regular smoking
- Recent smoking

Control: Background

- Gender
- Race
- Age
- Parent education
- Years residence in area
- Father presence

Control: Parent and Friend Smoking

- Mother approval of adolescent smoking
- Smoking by mother
- Smoking by three best friends
- Smoking by best friend

Control: General Social and Psychological

- Popularity
- Rejection
- Egotism
- Social passivity
- Reputation
- Femininity
- Self-deprecation
- Depression
- Boredom
- Sensation-seeking
- Deviance
- Rebelliousness
- Religious involvement
- Risk-seeking
- Locus of control
- Impulsivity
- Closeness to mother
- Peer versus parent influence

Control: Media Use

- Number of hours listen to radio
- Number of hours watch television

Control: Exposure to Antismoking Interventions

- Advertisements
- School curricula

Control: Other

- Smokeless tobacco use
- Alcohol use*

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*Variable used only in analyses to determine if campaign effects varied across subgroups.

A Comparison of Injury Death Rates in China and the United States, 1986

Guohua Li, MB, MPH, and Susan P. Baker, MPH

ABSTRACT

Background: The importance of injury as a public health problem is not well recognized in many developing countries. Data have recently become available on injury mortality in China.

Method: We compared Chinese injury data based on a 100 million population segment for 1986 with data for the United States.

Results: The age-adjusted death rate from all injuries for China exceeds the rate for the US (69.0 vs 61.3 per 100,000). The US has higher death rates from motor vehicle crashes, fires, and homicide; China has greater mortality from drowning, poisoning, falls, and suicide. Especially noteworthy in China are the high drowning rates among young children and the elderly and the high suicide rates in rural areas among young adults and the elderly.

Conclusion: Injury is an important public health problem in China, exceeding in many respects the problem in the United States. It is urgent for China to place high priority on injury research and prevention. (*Am J Public Health* 1991;81:605–609)

Introduction

Injury is now well recognized in the United States and other developed countries as an important public health problem, the leading cause of death from the first year of life to age 44.[1] Most efforts and achievements are limited to the US and other industrial nations. In developing countries, injury research is still in its infancy.

The People's Republic of China, with about one-fifth of the world's population, reflects many aspects of other developing countries. The authors compared the injury death rates in China and the United States to determine whether the great differences between China and the US in their social, economic, and cultural background are associated with major disparities in injury death rates.

Methods

China does not have a complete nationwide disease and injury surveillance system. Data on mortality have become available since the early 1980s with the development and extension of the disease surveillance points (DSP) in its 30 provinces, covering about 100 million population or about 10 percent of the total population of the country. The 1986 DSP injury data for China were provided by the World Bank and are published in summary form.[2] Data for the US come from the National Center for Health Statistics (NCHS).[3] Both sets of data are based on the Ninth Revision of the International Classification of Diseases.

Results

Crude and Age-adjusted Death Rates

The US has a higher crude death rate than China for all injuries combined, but the Chinese rate is higher when adjusted to the 1986 US population (Table 1). The death rates from falls and suicide for China increase significantly after adjustment, as rates from these causes are very high for the elderly and the American population is older than China's. Chinese age-adjusted death rates from falls, drowning, and suicide are two to four times the US rates.

For all injuries, the US demonstrates much greater gender differences. In US males, the age-adjusted rate is 2.7 times the rate for females. In China, the male rate is only 1.1 times the female rate.

Percent of Injuries by Cause

The top three causes of injury death in China are suicide, motor vehicle crashes, and drowning, which account for 33 percent, 16 percent, and 14 percent of all injury deaths, respectively (Figure 1). Motor vehicle crashes, suicide, and homicide are the leading causes of injury death in the US, accounting for 31 percent, 20 percent, and 14 percent of all injury deaths, respectively. The prominence of falls, poisoning, and suffocation is similar in the two countries.

Death Rates by Cause and Area

In both China and the US, urban areas have higher death rates from poisoning and falls while rural areas have higher death rates from fires/burns and drowning. For these causes, the urban/rural pattern differs markedly between the coun-

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From the Injury Prevention Center, Johns Hopkins School of Hygiene and Public Health. Address reprint requests to Guohua Li, MB, MPH, Injury Prevention Center, Johns Hopkins School of Hygiene and Public Health, 624 N. Broadway, Baltimore, MD 21205. This paper, submitted to the Journal May 31, 1990, was revised and accepted for publication October 10, 1990.

TABLE 1—Adjusted Death Rates* from Injuries by Cause and Sex in China and US, 1986

	Male		Female		Total		
					China		
Cause	China	US	China	US	Crude	Adjusted	US
Motor Vehicle	12.7	28.4	5.9	11.0	8.5	9.5	19.4
Poisoning	3.7	3.5	2.8	1.3	2.6	3.3	2.4
Falls	9.2	5.1	14.7	4.4	5.8	12.0	4.7
Fires and Burns	1.6	2.5	1.7	1.5	1.0	1.7	2.0
Drowning	10.1	3.3	6.0	0.7	7.7	8.1	2.0
Suffocation	1.9	1.7	1.7	1.4	1.7	1.8	1.5
Suicide	17.7	20.6	24.3	5.4	17.6	21.1	12.8
Homicide	1.9	14.1	0.9	4.2	1.4	1.4	9.0
All Injuries	71.4	90.6	65.6	33.6	54.1	69.0	61.3

*Adjusted to the 1986 US population.

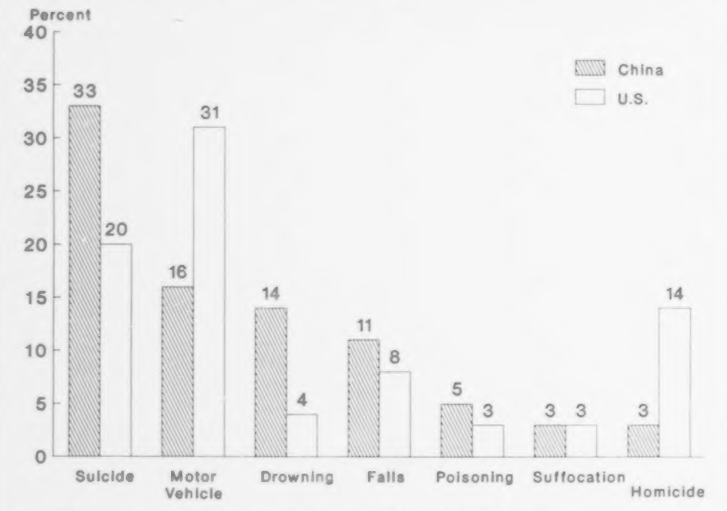

FIGURE 1—Percent of Injury Deaths by Cause in China and US, 1986

TABLE 2—Crude Death Rates from Injury by Cause and Area, in China and US*

	China (1986)		US (1980–1986)	
Cause	Urban	Rural	Central Cities	Remote Rural
Motor Vehicle	8.4	8.6	17.0	29.2
Poisoning	2.7	2.4	3.0	1.9
Falls	6.7	4.6	6.6	3.9
Fires/Burns	0.7	1.4	2.6	3.6
Drowning	4.0	12.7	2.0	5.9
Suffocation	0.6	3.1	0.4	0.6
Suicide	10.0	27.7	13.6	12.7
Homicide	1.6	1.2	17.7	5.4
All Injuries	41.4	71.0	73.4	80.7

*Source: Baker SP, et al.[10]

tries. In the US, homicide rates are very high in large cities and motor vehicle death rates are high in rural areas. Suicide rates, on the other hand, are similar in urban and rural areas of the US, while in China the rates in rural areas are almost three times the urban rates (Table 2).

Death Rates by Age and Sex

All Injuries: China has higher death rates at ages 0–14 and above 55 years, while the US has higher death rates at ages 15–54 (Table 3). The greatest disparity is at the two extremes of age; for children under one year the death rate in China is almost five times the US rate, and for age 85 and older, the rate is 2.5 times the US rate. For both countries there is a peak at ages 20–24 and a sharp rise after age 65. For males, the peak at ages 20–24 is more extreme in the US, where death rates exceed the Chinese rates from ages 15 to 59. In contrast, death rates for Chinese females are consistently higher than those for US females, except at ages 25–29. There is a pronounced peak in the death rate at ages 20–24 for Chinese females, but not for US females.

Motor Vehicle Crashes: Motor vehicle death rates in the US are more than double the rates in China for both males and females ages 10 to 29 (Figure 2). A peak at ages 15–24 characterizes the American death rates for both sexes. Otherwise, death rates increase fairly steadily throughout the life span in both countries.

Poisoning: High-risk populations for poisoning in China are infants (under one year) and the elderly (above 60 years), and the peak appears at age 85 and over for both males and females. In the US, the death rate also increases in the elderly, but is highest in males in their thirties.

Falls: In both China and the US, the high-risk population for falls is the elderly. Falls are the leading cause of fatal injury among ages 65 and above for both males and females in the two countries. The death rates for males at age 85 and above are 301.7 and 163.1 per 100,000 in China and the US, respectively. For females, the corresponding figures are 404.4 and 121.8. The Chinese rates are more than four times the US rates for children less than 15 years old.

Fires and Burns: Americans have higher death rates from fires and burns before age 70 than Chinese. However, Chinese rates are higher after age 70 than American rates. Notably, the US death rate from fires and burns under age five is more than twice the Chinese rate for both males and females.

TABLE 3—Death Rates from All Injuries by Age and Sex, in China and US, 1986

		<1	1–	5–	10–	15–	20–	25–	30–	35–	45–	55–	65–	75–	85+
Male	China	141	63	36	24	42	68	44	54	58	56	79	143	285	671
	US	33	27	16	23	102	141	121	107	93	87	88	110	209	422
Female	China	141	48	15	12	39	57	29	32	33	37	56	108	277	692
	US	30	19	9	10	34	34	30	27	26	28	32	46	95	221

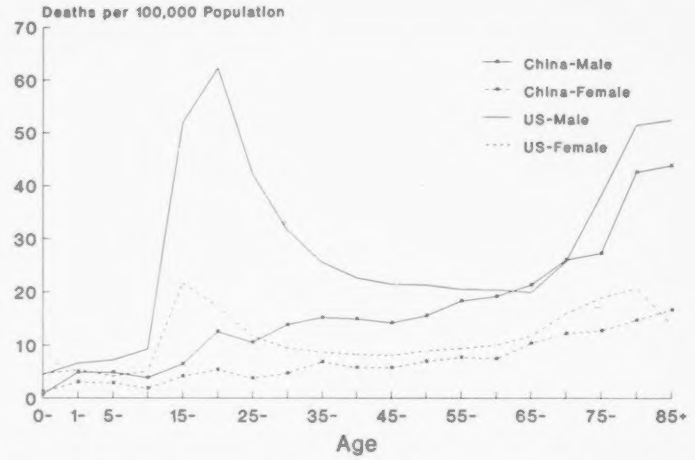

FIGURE 2—Death Rates from Motor Vehicles by Age and Sex in China and US, 1986

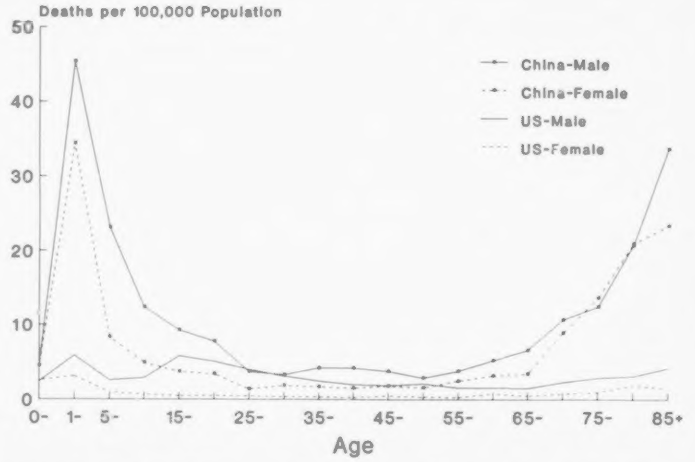

FIGURE 3—Death Rates from Drowning by Age and Sex in China and US, 1986

Drowning: Drowning is the third leading cause of injury death in China. The drowning rates at ages 1–4 for China are 45.4 and 34.4 deaths per 100,000 for males and females, respectively (Figure 3), seven to 10 times the American rates. Except for males ages 25–29, China has higher death rates from drowning for both males and females than the US.

Suffocation: Suffocation accounts for about 3 percent of all injury deaths. In infants, the rates in China are 19 times and 31 times the US death rates for males and females, respectively. Above age 85, on the other hand, the US rates are 42 times and 12 times the Chinese rates for males and females, respectively. In most other age groups the two countries share very low death rates from suffocation.

Suicide: Suicide is the leading cause of injury death in China and ranks second in the US. For males, China has lower death rates before age 60 and higher rates after age 60 than the US (Figure 4). For Chinese females, the suicide rate is more than 10 times the US rate at ages 15–24 and after 60, and more than twice the US rate for ages 25–59.

Homicide: In contrast to suicide, homicide is a major killer only for Americans. For each sex and age group, the US has higher homicide mortality (Figure 5). In both countries, male death rates are two to three times the rates for females. For the US, homicide rates are highest in the twenties and then decline until the sixties. For children less than 5 years old, the homicide rate in the US is more than four times the rate in China.

Discussion

Injury mortality data for China have not previously been described in detail. The general profile of injury death rates in China differs profoundly from the US. China's age-adjusted mortality from all injuries surpasses that of the US, even though its crude rate is lower than the US rate. In particular, China has higher death rates from falls, drowning, and suicide, and the US has higher rates from motor vehicles, fires and burns, and homicide.

Although the population-based death rate from motor vehicles in the US is twice as high as in China, there is evidence that China does not have safer transportation than the US. On average, China has about 15 motor vehicles per 1,000 population, which is less than a fortieth of the US ratio.[4,5] Its death rate per 100,000 motor vehicles is more than 10 times the rate of the US. The mortality from motor vehicles in China has been rising in the last 10 years[4] while declining in the US and other industrial nations.[6,7] With industrialization of China, death rates will continue to rise and the tragedy experienced by developed countries will be repeated, unless China pays enough attention to highway safety research and learns lessons from developed nations.

The very high suicide rate in China observed among ages 15–24, and among females age 60 or older, reflects many so-

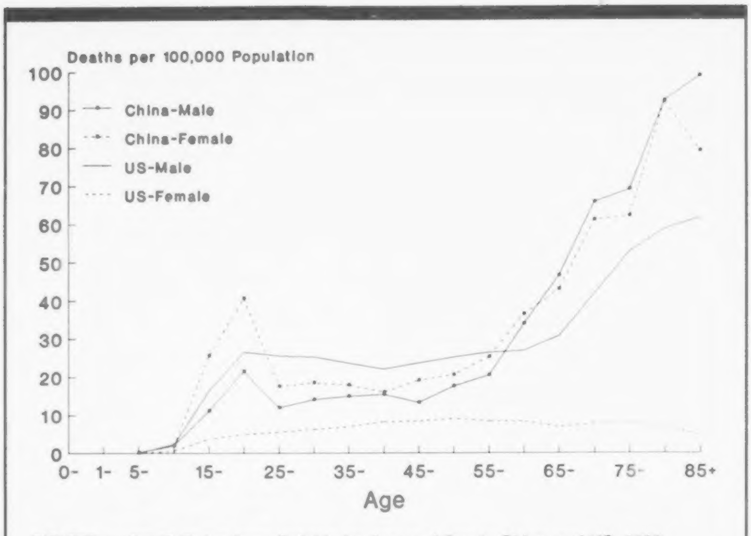

FIGURE 4—Death Rates from Suicide by Age and Sex in China and US, 1986

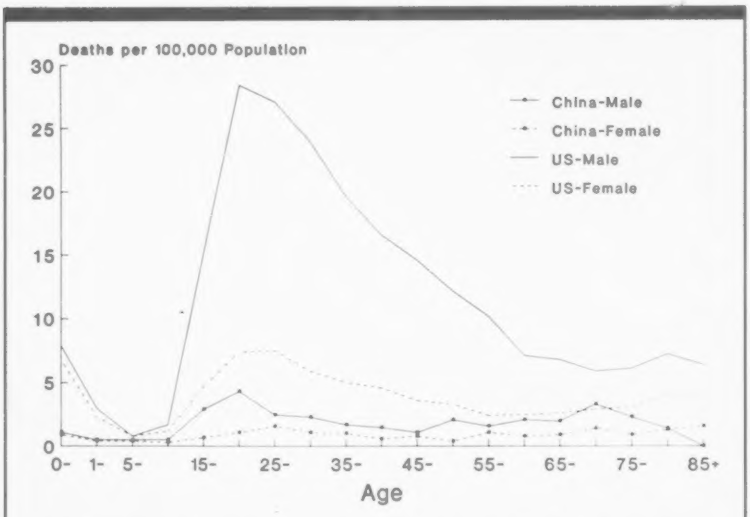

FIGURE 5—Death Rates from Homicide by Age and Sex in China and US, 1986

cial and cultural problems that China is facing. The motivation and cause of suicide may be greatly different between Chinese and Americans. For young Chinese females, especially those living in rural areas, marriage problems and poverty may be the major causes of the high suicide rate. For Chinese females at ages 20–24, the suicide rate is as high as the motor vehicle death rate for the American males in the same age group. In this age group, the suicide rate among females in rural areas is five times the rate in urban areas (78.3 vs 15.9/100,000). Similarly, for Chinese males ages 20–24, rural areas have four times the suicide rate as urban areas (40.7 vs 9.9/100,000). For Chinese elderly, poverty, long-term disease, and the decrease of traditional family relationships may contribute to the extremely high suicide rates. In the US, on the other hand, the suicide rate is highest in large cities and lowest in very rural areas. Because of the very complicated factors influencing suicide, China, as other countries, may have even greater difficulty in reducing its high suicide rates than in controlling its rising mortality from motor vehicles.

In contrast to suicide, the US has much higher homicide rates than China for each sex. Firearms are the weapon most commonly used in American homicide.[8,9] The law prohibiting individual ownership of firearms and virtual eradication of drug abuse may be major contributors to the relatively low homicide rate in China. On the other hand, even the non-firearm homicide rate in the US is substantially higher than China's total homicide rate, suggesting the important contribution of other factors.

Drowning is the third leading cause of injury death in China. Its major victims are young children and the elderly. Drowning rates in the US are highest at ages 1–3,[10] but are much lower than in China. One of the most effective interventions is barriers for swimming pools, which could reduce about 70 percent of childhood drownings in pools.[11] As drowning occurs in ponds, lakes, rivers, and other natural settings in China, to prevent and reduce the incidence is very difficult. In some rural areas of south China, drowning is the leading killer of children at ages 1–7.[12]

Among all injuries, falls exhibit the most similar profile in the two countries, although China has higher death rates at each age. Studies show that effective measures can reduce the incidence for children.[13,14] The higher death rates in China from falls in the elderly may reflect poor access to medical care, since mortality from falls often results from complications following relatively minor injury and has been greatly reduced in the US in recent years.[10]

Although the injury data for China are reasonably valid, some problems affect our comparison. First, about 80 percent of the Chinese population live in rural areas, but the rural population in the disease surveillance points sample accounts for only 43 percent. This should greatly underestimate the total death rate from all injuries in China, since the rate is much higher in rural areas than in urban areas. In particular, it leads to an underestimation of China's suicide and drowning rates, which are especially high in rural areas. Second, as ICD-9 has not been widely used in China, some misclassifications probably exist, and it is difficult to determine the degree of misclassification. Even so, our results suggest that injury is a serious public health problem in China that merits high priority. Special studies and injury surveillance systems that will provide details on the circumstances of injury are needed, followed as soon as possible by implementation of programs for effective injury prevention. □

Acknowledgments

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Core Curriculum on Tuberculosis Available through CDC

A training document on tuberculosis is now available through the Centers for Disease Control (CDC). The document was developed as part of the National Tuberculosis Training Initiative (NTTI) which was implemented in 1989 by the Division of Tuberculosis Control of CDC, the American Thoracic Society, and over 22 national professional medical and nursing associations including the American Public Health Association. Increasing the knowledge of health care providers about tuberculosis is a critical element in the Strategic Plan for the Elimination of Tuberculosis in the United States, which was recently endorsed by APHA.

The "Core Curriculum on Tuberculosis" contains information on epidemiology, transmission and pathogenesis, screening and prevention, diagnosis, treatment, infection control, and public health aspects of tuberculosis. It includes information on managing patients with coexisting human immuno-deficiency virus (HIV) infection. It is intended for use in developing educational programs as well as serving as a reference manual for health care providers caring for persons with or at risk of clinically active tuberculosis and tuberculous infection.

Copies may be obtained by writing to: Ms. Landis Brown, Technical Information Services, Center for Prevention Services, Centers for Disease Control, Mailstop E-06, Atlanta, GA 30333. Members of the American Public Health Association should identify themselves as such.

A slide series has been developed to accompany the "Core Curriculum on Tuberculosis." This series may be obtained on loan (and duplicated) from your State tuberculosis control program.

Comparison of Two Growth Charts in Lesotho: Health Workers' Ability to Understand and Use Them for Action

ABSTRACT

Background: Two growth monitoring charts widely used for growth monitoring in Africa (the Road-to-Health (RTH) and the Growth Surveillance (GS)) were compared in order to assist the Government of Lesotho to decide on an appropriate national growth chart.

Methods: Thirty-four health workers were taught and tested on the RTH during a first week of training and on the GS during a second week (the RTH-GS group), while the order was reversed for another 25 trainees (the GS-RTH group). The health workers were trained and tested on their ability not only to use and interpret the two charts, but also to make the right decisions about specific actions to be taken when growth faltering occurs.

Results: There was no difference between scores to the RTH and GS charts after one week of training. After the second week of training, the scores to the RTH chart improved and became better than those to the GS chart. The scores to the GS test did not increase with previous knowledge of the RTH chart.

Conclusions: For this reason and others discussed in the paper, the RTH chart was recommended for nationwide use in Lesotho. The adoption of this recommendation was facilitated by the close involvement in this research of public and private agencies responsible for growth monitoring in Lesotho. (*Am J Public Health* 1991;81:610–616)

Marie T. Ruel, PhD, David L. Pelletier, PhD, Jean-Pierre Habicht, PhD, John B. Mason, PhD, C. S. Chobokoane, PhD, and Arapang P. Maruping, MD, MPH

Introduction

Growth monitoring has been promoted by international agencies and local governments as one of the principal activities to combat malnutrition among young children in developing countries.[1,2] Growth charts are used

- as an educational and promotional tool, i.e., to make the dynamic process of growth visible to both health workers and mothers and to motivate them to take action to improve or maintain the child's nutritional status[3] and
- as a screening tool to allow the early detection of growth faltering in children and the targeting of appropriate interventions.[4]

Four conditions are necessary for growth charts to be useful for either educational or screening purposes. First, the child's age and weight data must be reasonably accurate. Second, health workers must be able to understand and interpret growth charts properly. Third, once growth faltering is detected, health workers must be able to identify the appropriate actions that need to be taken. Fourth, mothers must also be able to understand the growth charts and the messages provided by the health workers. The second and third issues are addressed in the present paper and the fourth one is discussed elsewhere.[5]

In Lesotho, a small kingdom surrounded by the Republic of South Africa, the concern about health workers' and mothers' ability to understand growth charts originated from the widespread use of two different charts throughout the country. The Ministry of Health, which administered approximately one-third of the Primary Health Care clinics, was using the World Health Organization (WHO)[6] version of the Road-to-Health chart (Figure 1). The remaining clinics were administered by the Catholic Relief Services and used the Growth Surveillance system.[7] This system consists of two charts: the Master Chart (Figure 2A), retained at the clinic for statistics compilations and the Growth Surveillance chart (Figure 2B), taken home by mothers.

In Lesotho, both charts had strong proponents. Those advocating the Growth Surveillance chart emphasized the fact that this chart had been developed in Africa and was more socially and culturally acceptable for African health workers and mothers. Although the superiority of this chart over the Road-to-Health had never been tested, it was assumed that, for Africans, a straight flat line would be a better graphical representation of adequate growth (regular weight increase) than an ascending slope. On the other hand, the proponents of the Road-to-Health chart felt that African mothers were not different from other mothers in understanding a positive slope as representing adequate growth and a negative slope as a sign of weight loss. Furthermore, proponents of the Road-to-Health chart felt that it was easier for health workers to learn and

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Address reprint requests to Marie T. Ruel, Institute of Nutrition of Central America and Panama (INCAP), Carretera Roosevelt, Zone 11, Apartado 1188, Guatemala, Guatemala, Central America. Drs. Pelletier and Habicht are with the Division of Nutritional Sciences, Cornell University, Ithaca, NY; Dr. Mason is Technical Secretary, ACC/SCN, World Health Organization, Geneva; Dr. Chobokoane is with the Food and Nutrition Coordinating Office, Prime Minister's Office, Maseru, Lesoto; Dr. Maruping is WHO Medical Officer, Harare, Zimbabwe. This paper, submitted to the Journal March 13, 1990, was revised and accepted for publication December 17, 1990.

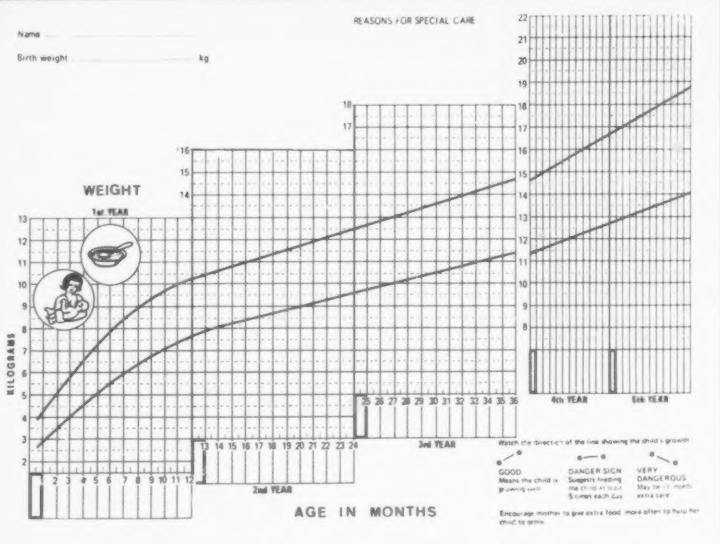

FIGURE 1—The Road-to-Health (RTH) Chart

interpret than the Growth Surveillance chart. UNICEF fostered this study to resolve these differences of opinion.

A Steering Committee, composed of members from various institutions involved in growth monitoring in Lesotho (see Appendix I for the list of participating institutions), was formed to advise the Government of Lesotho in the choice of a single growth chart to be used uniformly throughout the country. The final decision was to be based on the result of two studies, one that assessed health workers' ability to understand and use the charts for decision making (present study) and one that compared mothers' understanding of the two charts.[5] This Committee was involved in discussions and decisions about the various aspects of the research, including setting the objectives, choosing the experimental design, and developing the questionnaires and testing instruments. It also participated in the development of "Guidelines for Action", i.e. guidelines for decisions to be taken from the results of a weighing (Appendix II). Finally, the Steering Committee played a crucial role in facilitating implementation of the study in the field and in reviewing the results and discussing the inferences and implications for the country.

The objectives of the present study were to evaluate health workers' understanding and skills in using and interpreting the two growth charts after one week of training in each one, and in using the information for counseling and referral decisions according to the actions the Committee felt were indicated.

Methods

Study Sample and Design

Fifty-nine health workers from four hospital training centers of Lesotho were included in the study. Each participant was taught the two charts on two consecutive weeks.

A cross-over design was used to examine and take into account possible carry-over effects occurring when the learning of a chart taught second is made easier by the knowledge of the first chart. Therefore, two different treatment sequences were used and training centers were randomly assigned to either one: Hospitals A and B were taught the Growth Surveillance chart first and the Road-to-Health chart second (GS-RTH group), while Hospitals C and D were reversed and taught the Road-to-Health chart first and the Growth Surveillance chart second (RTH-GS group).

Growth Charts Used

The two charts compared were the RTH (Figure 1) and the GS (Figure 2B) chart. On the RTH chart, weight in kilograms forms the vertical axis and age of the child in months (ranging from 0–60 months) is on the horizontal axis. Two curves are printed on the chart and delimit the "Road-to-Health" zone. The upper one represents the median values for the reference population (50th percentile of the National Center for Health Statistics standards for boys[6]) and the lower one represents the NCHS third percentile for girls.[6] The GS chart contains the *percentage* of median weight-for-age (using the Harvard standards)[7] on the vertical axis and age in months on the horizontal one. Thus a straight line on this chart represents a child whose growth continues at a constant percentage of the standard weight-for-age. A green horizontal band (black in Figure 2B) delimits the so-called zone of "adequate growth" which corresponds approximately to the "Road-to-Health" on the RTH chart. A Master Chart (Figure 2A) must be used before plotting the data on the GS chart to find the corresponding percentage of standard weight-for-age for each child.

The decisions to be taken after weighing a child as developed by the Steering Committee (Appendix II) were also taught and used in the final test to assess the health workers' ability to use the charts for decision-making.

Teaching and Testing Procedures

One experienced teacher, who had been trained by the Catholic Relief Service and the Ministry of Health staff on the GS and the RTH charts, respectively, conducted all the teaching sessions. The format of the lecture on each chart had been standardized so that the amount of time spent on each aspect of the chart was the same for both. All lectures were taped and reviewed by the project director (MR) to ensure uniformity of content and lack of bias in the teaching.

On their first day, all participants were tested on their knowledge and understanding of both charts (baseline test) and information was collected on their previous exposure to the charts and work experience in primary care clinics. Each week of training in each chart consisted of approximately six hours of theory and six hours of practice. The training was done in a classroom setting using case histories. The trainees were instructed on how to calculate age given the date of birth (using a standard technique), how to plot hypothetical weights on growth charts, how to interpret growth curves, and what decisions to make (see Appendix II: Guidelines for Action).

The participants' knowledge, interpretation of the chart, and ability to make the appropriate decisions as prescribed by the Guidelines for Action were tested at the end of each training week: The final tests (one for the RTH and one for the GS)

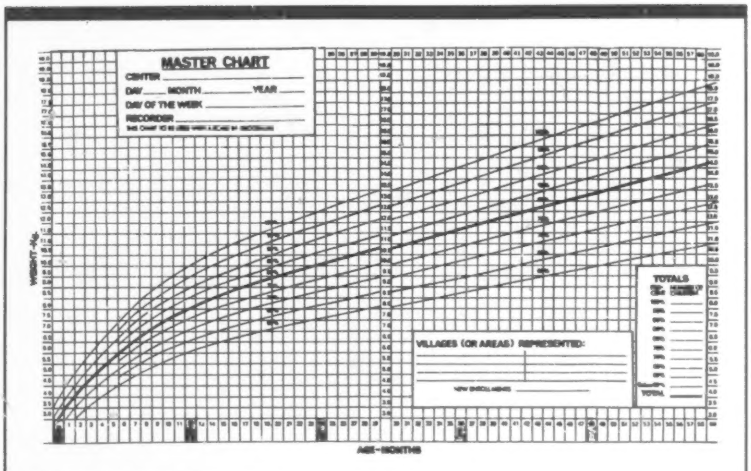

FIGURE 2A—The Master Chart

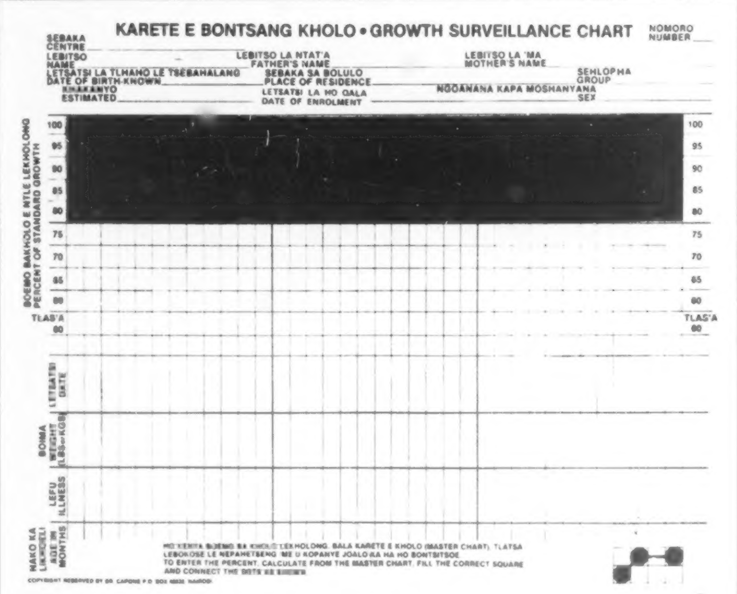

FIGURE 2B—The Growth Surveillance (GS) Chart

were very similar in their format and degree of difficulty. They were composed of case histories with questions on: age calculation, plotting of weights on the charts, interpretation of growth charts, and ability to use the Guidelines for Action.

Analytical Methodology

The total scores to the final tests as well as their different components were compared using t-tests, and confidence intervals were calculated using standard techniques.[8] The multivariate analysis of variance procedure (MANOVA)[9] was used to test main effects (chart, teaching period, training center) and interactions (particularly the carry-over effect, or treatment by period interaction); to control for baseline scores, age calculation scores and inter-hospital variability (using a nested design); and to obtain adjusted means and standard errors. The data were coded, entered and analyzed in Lesotho, using the Lesotho Food and Nutrition Coordinating Office (FNCO) microcomputer facilities, and the SPSS/PC+ statistical package.

Results

The overall score to the baseline test on growth charts knowledge was 3.4 out of 10 possible points and did not differ between training centers. Although work experience varied between participants (from 0 to 48 weeks), their poor knowledge of growth charts suggests that they had never received proper training in either chart.

A positive carry-over effect was found in the ANOVA model (significant interaction between chart and teaching schedule), meaning that prior training in one chart facilitated the learning of the second chart. Due to these findings, comparisons between the charts are presented separately for first training and for second week training.

Table 1 shows the results of the tests after training for all participants grouped by hospital and training schedule. In spite of similar baseline results, some differences between training centers were observed in the final scores. On average, Hospital B had higher scores than the other training centers, for both final tests, while Hospital C had lower scores. To control for the hospital variability, further analyses were done by analysis of variance (MANOVA procedure[9]), using a nested effect model (hospital within training schedule group). The baseline scores were also included in the model as covariates to adjust the final scores. The results of this analysis are presented in Table 2, which shows a comparison of the RTH and GS scores (unadjusted and adjusted) obtained on first week training and upon second week training, for the four hospitals grouped by training schedule. On first week training, the difference between the charts was minimal. Upon second week training, however, the scores to the RTH test were 15–20 percent higher (adjusted and unadjusted means, respectively) than the scores to the GS chart. This large difference between the charts resulted from a positive carry-over effect from first to second week training, observed only when the RTH chart was taught second: the GS-RTH group improved their score markedly from the first to the second week of training (+5.82, unadjusted means), while the RTH-GS group decreased their score from the first to the second week of training (−.76, unadjusted means). Examination of the components making up the total score to the final tests reveals that most of the carry-over effect seen in the GS-RTH group was due to the marked improvements in the interpretation and action questions (Table 1, Hospitals A and B).

TABLE 1—Results of the Final GS and RTH Tests, by Hospital and Training Schedule[a] (Group)

	GS Test				RTH Test			
Group/Hospital	Age[b]	Plotting[c]	Interpretation/ Action[d]	Total	Age	Plotting	Interpretation/ Action	Total
Group 1: GS-RTH[e]								
Hospital A	4.54	3.45	25.15	33.16	4.46	4.46	31.92	40.84
	(1.13)	(1.27)	(8.63)	(9.67)	(1.39)	(0.88)	(5.63)	(6.52)
Hospital B	4.00	4.00	33.90	41.90	4.09	4.38	38.09	46.57
	(0.95)	(1.00)	(6.47)	(7.30)	(1.73)	(1.28)	(3.05)	(4.33)
A + B	4.20	3.79	30.56	38.56	4.23	4.41	35.73	44.38
Pooled	(1.04)	(1.12)	(8.43)	(9.22)	(1.60)	(1.13)	(5.14)	(5.90)
Group 2: RTH-GS[e]								
Hospital C	3.10	2.30	23.50	28.90	3.50	3.20	24.00	30.70
	(1.52)	(1.70)	(8.42)	(11.03)	(0.97)	(1.50)	(8.31)	(9.47)
Hospital D	3.67	2.80	30.20	36.67	3.73	4.06	28.93	36.72
	(1.59)	(1.57)	(9.06)	(11.07)	(1.62)	(1.28)	(8.80)	(10.38)
C + D	3.44	2.60	27.52	33.56	3.67	3.72	26.96	34.42
Pooled	(1.56)	(1.61)	(9.26)	(11.50)	(1.38)	(1.43)	(8.79)	(11.50)

[a] The sample sizes were: Hospital A (n = 13), Hospital B (n = 21), Hospital C (n = 10), Hospital D (n = 15).
[b] Age: Calculation of age (maximum score = 5 points). Method was the same for both charts.
[c] Plot: Plotting of weights on charts (maximum score = 5 points).
[d] Interpretation and action: Interpretation of growth patterns and decision about actions to be taken, using the chart-specific Tables of Action (maximum score = 41).
[e] GS-RTH: GS taught first; RTH-GS: RTH taught first.
NOTE: The numbers in this table are means with standard deviations in parentheses.

Age calculation scores, on the other hand, did not improve from first to second training, but differed between hospitals, following a similar pattern as that observed for the overall scores. Since age calculation was independent of the charts, the differences observed between hospitals reflected variations in intrinsic learning capabilities of the health workers. An analysis of the sum of the interpretation and action scores was done that controlled for both hospital variability (using the nested design described previously) and differences in learning capability as proxied by age calculation confirmed the inferences drawn from Table 2. There was no substantial difference between the scores of the RTH and GS charts, −2.17 (with 95% confidence intervals (CI) of −6.67 to 2.33) after the first week of training. A significant improvement in the adjusted interpretation and action score from 31.81 ± 1.28 (standard error) to 38.16 ± 1.11 occurred when the RTH was taught in the second week. This contrasted with no improvement (from 33.98 ± 1.84 to 33.86 ± 1.59) when GS was taught in the second week. Thus the final adjusted score for interpretation and action was 4.3 (with CI of .43 to 8.17) in favor of RTH after the second week of training.

Discussion

In contrast to a previous study that tested health workers' skills in using different growth charts,[10] our study standardized the educational input across charts, and the participants were tested not only on their ability to understand, use, and interpret growth charts but also on their skills in using the Guidelines for Action to make the right decisions about a particular child. These guidelines were developed by Basotho experts involved and experienced in growth monitoring and maternal and child care. The feasibility in actual field conditions of implementing the actions proposed was an important consideration in the creation of these guidelines.

The present study had two major findings. First, at the end of the first train-

TABLE 2—Mean Scores to the GS and RTH Tests, by Group (Schedule of Training) and Week of Training[a]

	Results to the Final Test on		Difference between Charts
	GS Chart	RTH Chart	Difference (95% CI[b])
First week training			
	Group 1: GS-RTH (Hospitals A and B)	*Group 2: RTH-GS* (Hospitals C and D)	
Unadjusted means	38.56	34.32	+4.24 (−.84,9.32)
(SE)	(1.58)	(2.06)	
Adjusted means[c]	37.41	35.57	+1.84 (−3.94,7.63)
(SE)	(1.67)	(2.36)	
Second week training			
	Group 2: RTH-GS (Hospitals C and D)	*Group 1: GS-RTH* (Hospitals A and B)	
Unadjusted means	33.56	44.38	−10.82 (−5.90,−15.74)
(SE)	(2.30)	(1.01)	
Adjusted means[c]	36.04	43.77	−7.73 (−2.46,−13.00)
(SE)	(2.20)	(1.55)	

[a] Sample sizes were: group 1 (n = 34) and group 2 (n = 25).
[b] 95% confidence interval.
[c] The means were adjusted for hospital variability and for baseline scores, using the MANOVA[9] procedure and a nested design model (hospital within group).

ing week, the health workers who were taught the RTH understood their chart as well as those who were taught the GS chart. Thus, the argument that the GS was easier to learn for inexperienced African health workers was not substantiated. Second, when the participants were taught a second chart, those taught the RTH chart scored approximately 15 percent higher than those who were taught the GS chart. This improvement was not observed when the schedule of teaching was reversed: RTH first and GS second.

The improvement when RTH was taught second was achieved mainly through higher scores in the interpretation and action questions which measured the health workers' ability to identify appropriate interventions to be given to hypothetical children from the case histories. Since the guidelines were almost identical for the two charts,* it was expected that participants would master them better at the end of the second week, irrespective of the teaching sequence. However, our results showed no improvement on these questions and in the total score for those taught the GS chart on the second week (RTH-GS group). This finding suggests that prior knowledge of the RTH did not facilitate the learning of the GS chart and may even have confused the trainees because the visual presentation of growth in the GS chart is counterintuitive.

This finding is important since it is related to a situation that exists in many African countries: health workers that have been using a RTH-type chart for many years have to change to the GS system when the Catholic Relief Service Food and Nutrition Program is introduced; this requires intensive retraining of all the health workers. If the prior knowledge and experience with another chart (in this case the RTH) did facilitate the learning of the GS chart, the whole effort involved in the change of system would be reduced and potentially justifiable. However, the results of our study suggest that this was not the case in Lesotho.

Therefore, our conclusion from this study is that, in Lesotho, the RTH would be a better choice than the GS chart for two reasons. First, since health workers could learn and understand the RTH as well as the GS chart after the first week of teaching, it would be an appropriate chart for training inexperienced people. Second, and more important, in the context of Lesotho where retraining is a crucial issue, the RTH chart would be superior since second week training on the RTH improved the scores of those previously trained in the GS chart.

These recommendations are in accordance with those made in a concurrent study[5] showing that mothers who were taught the RTH learned significantly more about their chart than those who were taught the GS chart. Thus our overall conclusion from these two studies is that, in Lesotho, the RTH would be a more appropriate national chart than the GS. Nevertheless, it would be useful to replicate these results elsewhere in Africa where literacy of mothers and health workers is lower and more typical of much of Sub-saharan Africa. It would also be interesting to re-assess health workers' ability to use and interpret the two charts after use in growth monitoring activities.

We believe that the use of information from applied research, such as this, is fostered by the involvement of the national decision makers at all stages of the research process. In Lesotho, the decision was made to use a RTH-type chart on a nationwide basis, as recommended by our study. □

Acknowledgments

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APPENDIX I

List of Institutions Participating in the Steering Committee in Charge of the Study

Name of Institution

Bureau of Statistics of Lesotho
Catholic Relief Services
Food and Nutrition Coordinating Office
Joint Program in Nutritional Surveillance of the UNICEF (Eastern and Southern Africa Regional Office) and the Cornell Nutritional Surveillance Program, Cornell University
Ministry of Agriculture of Lesotho
Ministry of Health of Lesotho
Private Health Association of Lesotho
Red Cross
UNICEF, Lesotho Country Program
World Health Organization

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*The only difference was that for the RTH chart, some recommendations were made for children with constant weight (Appendix II). For the GS chart this was not possible because constant weight cannot be observed directly on the chart.

APPENDIX IIA— Guidelines for Action Based on the Interpretation of Growth Patterns Plotted on the Road-To-Health (RTH) or the Growth Surveillance (GS) Chart

Part 1. *For Children Re-attending the Clinic*

	Age 0–24 months		Age ≥25 months		
Direction of Growth	Gaining Weight	Constant/ Losing Weight	Gaining Weight	Constant Weight (only for RTH[a])	Losing Weight
Position of Dot					
A) Above 80%	• Regular plan of action[b] • Praise/ encourage	• Refer to nursing sister	• Praise/ encourage	• Regular plan of action (if constant weight for 3 months: refer to nursing sister	• Refer to nursing sister
B) Below 80%	• Regular plan of action • Home visit	• Refer to nursing sister • Home visit	• Encourage • Follow-up monthly	• Regular plan of action (if constant weight for 3 months: refer to nursing sister	• Refer to nursing sister

Part 2. *For New or Irregular Attenders*

	Age 0–24 months	Age ≥25 months
Position of Dot		
A) Above 80%	• Use "thinness chart"[c]	• Encourage monthly visits
B) Below 80%	• Use "thinness chart"	• Use "thinness chart"

Results of "Thinness Chart"

Green	Yellow	Red
• Praise/encourage • Encourage monthly attendance	• Refer to nursing sister	• Refer to nursing sister (very critical in children <24 months)

[a] On the GS chart, constant weight is not detected until the child's % weight-for-height (WTHT) drops. In this case, constant weight is detected on the GS chart by a downward slope, similar to the effect of a weight loss.

[b] See Appendix IIB.

[c] The "thinness chart" is a weight-for-height wall chart.[11] The interpretation of the color code is:
Green: WTHT between 70% and 80% of the NCHS standards
Yellow: WTHT between 80% and 90% of the NCHS standards
Red: WTHT between 90% and 110% of the NCHS standards.

APPENDIX IIB

Description of the Regular Plan of Action

REGULAR PLAN OF ACTION

1. Inform mother about her own child's growth (as seen on the growth chart). This should be done with all mothers.
2. Briefly interview mothers about their feeding practices. (Consider each age group: 0–3 months; 4–9 months; 9–12 months; over 1 year).
3. Give advice to mothers concerning specific feeding practices for each age group (taking mothers' reported practices into consideration). Suggested advice for each age group:

0–3 months	*4–9 months*
• Breastfeeding • Introduce liquids	• Breastfeeding • Introduce soft foods: cereals, fruits, vegetables, mashed/ pureed eggs
9–12 months	*>12 months*
• Breastfeeding • Continue soft foods • Introduce solids • Start meat/fish/ chicken/peas/ beans • Give frequent meals (>4/day)	• Breastfeeding • Adult foods: 3 food groups • Frequent meals (>4/day)

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Public Health Briefs

Evaluating Needle Exchange: Do Distributed Needles Come Back?

Joseph Guydish, PhD, MPH, George Clark, Delia Garcia, Moher Downing, Patricia Case, and James L. Sorensen, PhD

ABSTRACT

We employed capture-recapture methods as a strategy for evaluating needle exchange. Needles distributed by the exchange at two time periods were marked with color coded bands indicating the date and site of distribution. Half of the marked needles (2,068/4,239) returned within two weeks of distribution, and 61 percent (2,593/4,239) returned during the study period. The rate of return for stationary exchange sites (63 percent) was greater than that for roving/mobile sites (51 percent; $\chi^2 = 28.6$, p < .001). Of all needles returned, 87 percent (2,248/2,593) returned to the site of original distribution. (*Am J Public Health* 1991;81:617–619)

Introduction

Needle exchange programs can potentially slow the spread of the human immunodeficiency virus (HIV) among injection drug users (IDUs), their sexual partners, and their children. Needle exchange is controversial, however, in that it contradicts traditional drug treatment philosophy and removes at least one barrier to drug use. Reports from programs abroad suggest that needle exchange does not increase drug use or needle use, or encourage individuals to begin using nedles.[1–5] Whether these programs slow the spread of HIV remains unanswered. United States needle exchange programs now operate in several urban areas,[6–8] and there is a pressing need for evaluation. Strategies that do not rely on self-report can assist efforts to determine program effects.

One question is whether distributed needles return to the exchange. In the ideal program 100 percent of needles distributed would be returned in a short time, implying consistent use of new needles and a low likelihood of needle sharing in the exchange population. We employed capture-recapture methods to determine what proportion of distributed needles return to the exchange, how quickly needles return, and whether distributed needles travel between exchange sites.

The San Francisco exchange program (Prevention Point) trades needles on a 1:1 basis during regular hours on one night each week, and any patron may exchange up to 10 needles. There are two "roving" teams which travel on foot over a prescribed route, and two stationary teams. Together these sites exchanged an average of 2,806 needles per week during the study period.

Methods

Needles were identified for study purposes ("captured") with colored bands painted onto the barrel of each syringe. The color indicated both the site and date of distribution. Marked needles were distributed on two occasions. At Time 1, all needles distributed at Stationary Site A (n = 1,200) and Roving Site A (n = 204) were marked. At Time 2, nine weeks after the first distribution, marked needles were distributed at Stationary Site A (n = 1,300), Stationary Site B (n = 1,137), Roving Site A (n = 249), and Roving Site B (n = 149). These needles, marked with four colors different from those used at Time 1, comprised 95 percent (2,835/2,978) of all needles distributed at Time 2. After each distribution, all needles returned to any exchange site were inspected for a nine-week period, and banded needles (those "recaptured") were recorded.

This method is observational and unobtrusive. Exchange patrons were not interviewed, and no attempt was made to determine whether the original recipient of marked needles was also the person who returned them. Only needles collected by the exchange program were monitored, so no data are available concerning marked needles that did not return to the program. Prevention Point had distributed marked needles in conjunction

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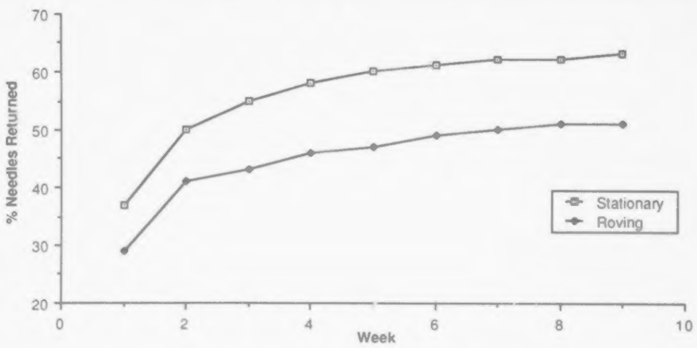
From the Center for AIDS Prevention Studies, University of California, San Francisco (Guydish, Sorensen) and the Prevention Point Research Group, San Francisco (all other authors). Address reprint requests to Joseph Guydish, PhD, Center for AIDS Prevention Studies, 74 New Montgomery Street, Suite 600, San Francisco, CA 94105. This paper, submitted to the Journal July 12, 1990, was revised and accepted for publication January 8, 1991.

TABLE 1—Rate of Return for Banded Needles, Collapsed across Time of Distribution and Site

Week of Return	Needles Returned N = 4,239* (%)
1	1500 (35.4**)
2	568 (13.4)
3	195 (4.6)
4	129 (3.0)
5	83 (2.0)
6	37 (0.9)
7	30 (0.7)
8	34 (0.8)
9	17 (0.4)
Totals	2593 (61.2)

*Total number of banded needles distributed.

**Percentages rounded to one decimal place.



Proportion returned for Stationary Sites is 62.8% (2284/3637), and for Roving Sites, 51.3% (309/602). $X^2 = 28.6$, $p < .001$

FIGURE 1—Cumulative Rate of Return by Type of Exchange Site, Collapsed across Time Periods

with a previous study,[6] so that many exchange patrons were accustomed to receiving marked needles. In the present study, patrons who asked about the markings were told that we wanted to learn how long distributed needles remained on the street.

Results

Results are presented for Time 1 and Time 2 combined, since the proportion of needles returned (60 percent and 62 percent, respectively) and the general pattern of return were the same. The rate of return was inversely related to time since distribution, with the greatest return occurring in the first five weeks (Table 1).

For the stationary site evaluated at both time periods, there was no difference in rate of return across time (63 percent vs 62.9 percent). For the roving site, the rate of return was greater at Time 2 than at Time 1 (58.2 percent vs 43.1 percent; $\chi^2 = 10.2$, $p < .01$). At Time 2 there were no differences in rate of return for the two stationary sites (62.9 percent vs 62.4 percent), or the two roving sites (58.2 percent vs 51 percent). To evaluate differences in return rate by type of site, we collapsed data across time for stationary and roving sites (Figure 1).

Last, we evaluated how often needles returned to the site where they were distributed. Of the banded needles returned, 87 percent (2,248/2,593) were returned to the site of original distribution. For stationary exchanges the proportion of needles returning to the home site ranged from 88 percent to 94 percent. The proportion returning home to Roving Site A was lower (74 percent). Unexpectedly, only 12 percent (9/76) of returning Roving Site B needles came back to the home site, and 82 percent returned to Stationary Site B located in the same neighborhood.

Discussion

For both Time 1 and Time 2, nearly 50 percent of distributed needles returned within two weeks. This finding suggests some stability in the exchange population, and that many patrons use the exchange repeatedly. The difference observed for rate of return between stationary and roving sites suggests that these models, in San Francisco, may serve different IDU populations. Roving teams were designed to reach homeless IDUs who are more transient and whose needles are frequently lost or stolen. To the degree that roving teams reach this population, a lower rate of return would be expected. The finding that needles distributed by Roving Site B frequently returned to Stationary Site B indicates that these sites serve an overlapping population.

The application of nonself-report methods is vital in adequately evaluating the effects of US needle exchange programs. While such methods cannot replace self-report strategies, they can provide independent verification of self-report data. The capture-recapture method indicates the stability of the exchange population, the frequency with which patrons use the exchange service, and the degree to which different exchange sites serve unique populations. ☐

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Targeted HIV Screening at a Los Angeles Prenatal/Family Planning Health Center

Laura J. Fehrs, MD, Deirdre Hill, MPH, Peter R. Kerndt, MD, Thanne P. Rose, PhD, and Carol Henneman, MD

ABSTRACT

In 1988, the Los Angeles County Health Department conducted a blinded human immunodeficiency virus (HIV) seroprevalence study at a public prenatal (PN) and family planning (FP) center serving mostly Hispanic women in order to determine seroprevalence and to evaluate the center's targeted HIV screening program. Four women (0.13 percent) tested positive (3/1801 PN and 1/1167 FP). Three reported no risk factors; one reported a history of syphilis since 1978. Voluntary HIV testing was selectively offered to women who reported risk factors for HIV infection. Only 14 percent (96/685) of clients offered testing chose to do it: 28 percent (14/50) of clients classified as being at highest risk of infection, and 27 percent (16/59) of women who judged themselves to have some chance of being exposed to HIV. None of the four women who tested positive by blinded testing chose testing. While few women at this center were infected with HIV, higher risk women were not persuaded to be tested through a targeted screening program. Blinded HIV seroprevalence studies provide a tool for both tracking infection in a population and evaluating screening programs. (*Am J Public Health* 1991;81:619–622)

Introduction

Women of childbearing age in the United States are at risk for human immunodeficiency virus (HIV) infection although their risk varies by locality, risk group, and race.[1–14] In 1987, the US Public Health Service (PHS) recommended that women with identifiable risks for HIV infection be routinely counseled and tested for HIV antibody.[15] Strategies for providing women of childbearing age with HIV counseling and testing include: 1) routine or universal screening, with informed consent, of women attending women's health clinics, such as prenatal (PN) or family planning (FP) clinics and 2) selective targeting of women identified as having risk factors for HIV infection. In the latter approach, risk factors for HIV infection are assessed during a counseling session and those patients reporting risk factors are offered voluntary testing. In areas with high rates of HIV seropositivity among women, this approach has not been shown to be effective in identifying most seropositive women.[3,4,8,10,16] Fewer data exist for areas with low or unknown rates of seropositivity among women, such as Los Angeles.

In 1988, the Los Angeles County Department of Health Services participated in the PHS Family of Surveys[17] and initiated a blinded HIV seroprevalence study at a large inner-city public health center which used a targeted approach for HIV counseling and testing in its PN and FP clinics. Results of blinded HIV testing were used to evaluate the extent to which the HIV counseling and testing program reached women who were either at risk for HIV infection or already infected.

Methods

Consecutive sera remaining after routine syphilis testing during initial or annual visits of all women attending the PN clinic from March through September, and the FP clinic from March through November, 1988 (3,000 total) were blinded by the removal of personal identifiers and then tested for HIV antibody by a licensed enzyme linked immunosorbent assay (ELISA) kit. Repeatedly reactive specimens were confirmed with immunofluorescent antibody (IFA) (Laboratory Services Branch, California State Department of Health Services) and Western blot.

The clinics used a targeted HIV counseling and testing program. All patients were counseled by trained bilingual interviewers about risk factors for HIV infection and the availability of HIV testing. The interviewers used a questionnaire

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Address reprint requests to Laura J. Fehrs, MD, Division of Immunization, Centers for Disease Control, Atlanta, GA 30333; Drs. Kerndt and Fehrs were assigned to the Los Angeles County Department of Health Services from the CDC Epidemiology Program Office at the time of the study. Ms. Hill and Dr. Kerndt are with the AIDS Epidemiology Office; Dr. Henneman is with the H. Claude Hudson Comprehensive Health Center. All are with the LA County Department of Health Services. Dr. Rose was with the AIDS Epidemiology Office at the time of the study. She is currently with the Academy of Pediatrics in Chicago, Ill. This paper, submitted to the Journal May 22, 1990, was revised and accepted for publication November 20, 1990.

TABLE 1—Definition of Risk for HIV Infection Used by Clinic Staff and by Investigators, and Prevalence of Risk Behaviors, Los Angeles, 1987–88

	Risk behaviors (+) for which clinic staff offered voluntary testing	Risk behaviors (+) for which women were classified as "high risk" by investigators	Women with risk behavior No. (%)
Intravenous (IV) Drug Use since 1978	+	+	8 (0.3)
Sex since 1978 with an IV drug user, bisexual man, or person with AIDS or HIV infection	+	+	47 (1.6)
Receipt of money or drugs for sex in the last year	+	+	12 (0.4)
Transfusion from 1978 through 1985	+		107 (3.6)
Sex since 1978 with a prison inmate	+		172 (5.7)
Sex since 1978 with a man with hemophilia	+		0 (0)
Sex since 1978 with a man from Haiti or Central Africa	+		6 (0.2)
History of hepatitis	+		48 (1.6)
History of a sexually transmitted disease since 1978	+		38 (1.3)
History of more than one sex partner in the last year	+		165 (5.5)
History of sexual assault	+		60 (2.0)

TABLE 2—Results of HIV Antibody Seroprevalence Study, Los Angeles, 1987–88

		HIV Antibody Seroprevalence		
	% of All Patients	No. HIV Positive	Total No. Tested	% Positive
Clinic Type				
Prenatal	60.0	3	1801	0.17
Family Planning	38.9	1	1167	0.09
Unknown	1.1	0	32	0
Risk For Infection				
No Risk Behaviors Reported	77.0	3	2311	0.13
"At Risk" (Any risk behaviors reported)	23.0	1	689	0.15
"High risk"*	1.7	0	50	0
Self-Assessed Risk of ≥ 1 in 100 chance**	2.0	0	60	0

*Subset of "At Risk" group; these two groups were not mutually exclusive. See Table 1 for definitions.

**Self-assessment of risk of exposure was determined by interviewer asking patient "How likely do you think it is that you have been exposed to someone infected with the AIDS virus in such a way that you could be infected?" Patients were given the choice of answering "Very unlikely (less than 1 in 100 chance)", "Unlikely (1 in 10 chance)", "Uncertain (50-50 chance)", or "Certain (sure that you have been exposed to someone infected with the AIDS virus)."

as a guide to counseling and as a record of risk behavior information. To identify all women possibly at risk, the clinic chose to use a broad range of behaviors in its definition of risk (Table 1) and patients were also asked to assess their own risk (Table 2). Voluntary HIV testing was recommended and immediately offered to all women who requested testing and to those reporting any of the behaviors defined by the clinic to indicate possible risk.

The targeted HIV counseling and testing program was evaluated by using a unique study number to link the results of blinded seroprevalence testing to questionnaire information. Women were classified by their answers as being "at risk," "at high risk," or "not at risk." The "at risk" group was composed of those women who reported any of the broad range of risk behaviors which the clinic considered to place women at some risk of infection (the group that was offered HIV testing). A subset of "at risk" women, "high-risk" women, was defined by investigators to be at particularly high risk for HIV infection, using objective criteria (Table 1).

Confidence limits for proportions were calculated using the method published by Fleiss.[18]

Results

A total of 1,801 consecutive specimens from the PN clinic, 1,167 consecutive specimens from the FP clinic, and 32 specimens for which clinic type was not known were tested for HIV antibody (Table 2). Of the patients for whom race was known (2,955), 87.3 percent (2,579) were Hispanic, 9.6 percent (285) White, 2.4 percent (71) Black, 0.5 percent (15) Asian, and 0.2 percent (5/2,955) other race. The median reported length of residence in the US was five years. The median age was 25 years with a range of 12–46 years; age was unknown for 29 women.

Four of the 3,000 specimens (0.13 percent, 95 percent confidence interval (CI) = 0.04, 0.37) were repeatedly reactive by ELISA and positive by IFA and Western Blot; three were from the PN clinic (0.17 percent, 95% CI = 0.04, 0.55), and one was from the FP clinic (0.09 percent, 95% CI = 0.00, 0.55) (Table 2). These positive specimens were from one White and three Hispanic patients, all of whom were in the 20–24 age group. The length of time that these patients reported residing in the US ranged from two to 12 years.

None of the four seropositive women chose voluntary HIV testing (Table 3). Only one of the four reported any risk factor for infection—a history of syphilis since 1978 (Table 2). Because this finding was of concern, we further investigated the center's testing strategy by determining the rate of acceptance of testing among those reporting risk behaviors.

While 23 percent of women reported one or more risk behaviors (were considered "at risk" by clinic staff) and 1.7 percent met the criteria for classification as "high risk," only 2.0 percent judged themselves to have more than a one in 100 chance of being exposed to HIV infection (Table 2). A high proportion of women who were "at risk" (92 percent) or even "high risk" (66 percent) assessed their own risk for being exposed to HIV infection to be "very unlikely" (less than a one in 100 chance) (Table 4). All four of the

TABLE 3—Acceptance of Voluntary HIV Testing, by Risk Group

Risk Group	No. Accepted Voluntary HIV Testing	Total No. in Group*	% Accepted Voluntary Testing
Reported Risk Behaviors			
"At Risk" (Any risk behaviors reported)	96	685	14.0
"High risk"	14	50	28.0
Self-Assessed Risk	16	59	27.1
Tested Positive in HIV Seroprevalence Study	0	4	0
Tested Negative in HIV Seroprevalence Study	110	2985	3.7

*Persons for whom acceptance of testing was not known were excluded from analysis

seropositive women also assessed their own risk as being "very unlikely."

Not only did the four seropositive women refuse HIV testing, but only 14.0 percent of women offered testing because of a history of risk behaviors ("at risk" group) consented (Table 3). Among the women who were classified as being "high risk" and among the women who judged their chance of exposure to be greater than one in 100, only 28.0 percent and 27.1 percent, respectively, selected voluntary testing.

Comments

The prevalence of HIV antibody among prenatal and family planning patients attending this inner-city, mostly Hispanic, public health center in Los Angeles was low—0.13 percent (95% CI = 0.04, 0.37). The rate found in this study is consistent with the low prevalence of antibody reported from parturients (0.04 percent),[19] newborn screening (0.10 percent),[20] and FP clinics (0.1 percent, D. Hill, unpublished data) in Los Angeles County. Since the population in this study was predominantly Hispanic and immigrant, it should not be considered representative of all clinics with a high proportion of minority clients. But, as one of six large comprehensive public health centers in Los Angeles, it is an important site for monitoring HIV infection.

While the seroprevalence found at this center was reassuringly low for an urban area with a high incidence of AIDS, it is of concern that the program of targeted HIV screening did not identify the seropositive women in this study. Three of these four women reported no risk factors for infection. The rates of reporting of risk behaviors and acceptance of testing in this population may have been affected by cultural factors and/or lack of knowledge of risk behaviors of partners. Some women may have been tested already for HIV infection for immigration or other purposes. Acceptance of testing may have been affected in unknown ways by the counseling process; however, in this regard, the clinic is probably typical of urban, high-volume clinics.

Our data suggest that a strategy of selectively offering HIV voluntary testing only to those women who report risk factors for infection may result in a low rate of acceptance and may miss seropositive women. Similar data from several studies led the US Public Health Service's Immunization Practices Advisory Committee to recommend the universal screening of PN patients for hepatitis B.[21–26] Universal screening for HIV of all women pregnant or contemplating pregnancy has been recommended[1–4,16,27,28] because it identifies a high proportion of infected women; these women can be referred early to medical prophylaxis and treatment; and it results in the early identification and treatment of children born to HIV-infected mothers. Universal screening, with informed consent, has been conducted at some Los Angeles prenatal clinics since 1988.

Despite the benefits of universal screening, targeted screening may be used in localities with a low prevalence of HIV or without resources for adding large scale HIV testing and counseling to routine clinical practice. Efforts to identify seropositive women compete for limited resources with HIV prevention and other public health programs. In a 1988–89 random survey of urban hospitals, most did not offer HIV testing routinely to pregnant women.[28]

Those who design voluntary HIV screening programs for women should consider: local seroprevalence; available resources vs demand for testing and other services; local characteristics of women at risk for infection; strategies that result in high levels of acceptance of testing among women at risk; and strategies that provide counseling to seronegative women and effect behavioral changes that will prevent the future transmission of HIV.

The variety of strategies for screening women of childbearing age should be evaluated locally and over time. Blinded HIV seroprevalence studies provide valuable data about voluntary screening programs. In addition, they supplement acquired immunodeficiency syndrome case

TABLE 4—Self-Assessment of Risk of Significant Exposure to HIV Infection, by Risk Group, Los Angeles, 1987–88

Self-assessment of Risk of Exposure	Risk Group					
	"High risk"		"At Risk"		No Risk Factors	
	No.	%	No.	%	No.	%
Very unlikely (<1 in 100 chance)	33	66.0	633	91.9	2268	98.1
Unlikely (1 in 10 chance)	6	12.0	17	2.5	14	0.6
Uncertain (50-50 chance)	10	20.0	19	2.8	6	0.3
Certain (sure was exposed)	1	2.0	3	0.4	1	0.04
No response	0	0	17	2.5	22	0.7
Total	50	100	689	100	2311	100

surveillance by tracking trends in HIV infection among women over time. ☐

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HIV-1 Prevalence in Selected Tijuana Sub-populations

Fernando Güereña-Burgueño, MD, Abram S. Benenson, MD, and Jaime Sepúlveda-Amor, MD

ABSTRACT

To assess the prevalence of HIV-1 (human immunodeficiency virus) infection among high-risk populations in Tijuana, Mexico, HIV-1 antibody status was determined and information on risk behavior was obtained from 1,069 individuals in three high-risk groups. The prevalence of HIV-1 among 415 prostitutes was 0.5 percent; 410 prisoners, 1.2 percent; 233 homosexual/bisexual men, 11.6 percent; and 106 intravenous drug abusers, 1.9 percent. The potential for spread of HIV-1 exists in Tijuana despite the current relatively low seroprevalence of HIV-1. (*Am J Public Health* 1991;81:623–625)

Introduction

By December 31, 1989, 3,512 AIDS (acquired immunodeficiency syndrome) cases have been reported to the Mexican Ministry of Health.[1] Projections made by the Mexican National Committee for AIDS Prevention (CONASIDA) suggest that the cumulative AIDS case total could exceed 75,000 by the end of 1991.[2] The majority of the 71 AIDS cases in the state of Baja California[2] have been reported by the city of Tijuana, geographically contiguous to San Diego, California. This border area is the busiest port of entry to the United States, with approximately 36 million legal entries of alien and US citizens, and more than 500,000 illegal aliens deported in 1987.

This study presents the results of a cross-sectional epidemiologic and serological survey designed to determine the prevalence of infection with the human immunodeficiency virus type 1 (HIV-1) among "high-risk" populations in Tijuana. The project was approved by the Committee on Protection of Human Subjects of the San Diego State University.

Methods

Sample sizes were selected to provide less than 3 percent margin of error ($\alpha = .05$ and $1\text{-}\beta = .80$) (4). From 1,967 prisoners in the Tijuana State Penitentiary, 410 were recruited by selecting every fifth prisoner listed in the prison records. From 2,000 active prostitutes registered at the Municipal Medical Services Clinic, 415 were sequentially included in the study. A convenience sample of 233 homosexual/bisexual men was recruited in diverse homosexual gathering sites. From 1,000 intravenous drug abusers with files at the Juvenile-Integration Center detoxification clinic at Tijuana, a sample size of 341 was included in the study by selecting every third individual listed in the clinic records.

Participation was voluntary and anonymous. Each subject was individually counseled prior to testing. Sociodemographic and risk factor information was obtained and a blood sample was drawn. Test results with post-test counseling were offered to the participants.

Antibodies against HIV-1 were determined with licensed enzyme-immunoassay (EIA) kits.[5] Those repeatedly reactive were then confirmed by the indirect immunofluorescent antibody assay (IFA)[6,7] or by the Western blot technique.[8]

Descriptive statistics and odds ratios were calculated and 95 percent confidence intervals estimated with Cornfield's method.[9]

Results

Questionnaires were completed and blood samples obtained from 1,069 individuals. The refusal rate was 2 percent in the prisoner and prostitute groups and 25 percent to 30 percent in the group of homosexual/bisexual men. Only 11 of the 341 drug abusers in the files of the detoxification clinic could be located. However, 95 individuals in the other three groups admitted intravenous drug abuse (IVDA).

A total of 34 individuals tested positive for HIV-1. The prevalence rates by study group, sex, and history of IVDA are shown in Table 1.

Seropositive homosexual/bisexual men exhibited high-risk behavior patterns (Table 2); a history of receptive anal sex showed the strongest relation to HIV-1 infection. For vaginal sex, 36 percent (139/383) of prostitutes reported using condoms "usually" or "always." Of female prostitutes and female prisoners participating in the study, 12.3 percent (46/383) and 5.9 percent (2/34), respectively, reported anal sex.

From the Graduate School of Public Health, San Diego State University California (Güereña-Burgueño and Benenson) and the General Directorate of Epidemiology, Ministry of Health, Mexico (Sepúlveda-Amor). Address reprint requests to Fernando Güereña-Burgueño, MD, Division of HIV/AIDS, Centers for Disease Control, 1600 Clifton Road, NE, Mailstop G-29, Atlanta, GA 30333. This paper, submitted to the Journal April 2, 1990, was revised and accepted for publication October 10, 1990.

TABLE 1—Prevalence of HIV-1 Antibody by Study Group, Tijuana HIV-1 Seroprevalence Study, June–October 1988

	Total		Intravenous Drug Abusers*		Non-intravenous Drug Abusers	
Study Group	No. Subjects	Prevalence (%)	No. Subjects	Prevalence (%)	No. Subjects	Prevalence (%)
Prostitutes						
Males	1/32	(3.1)	1/2	(50.0)	0/30	(0.0)
Females	1/383	(0.3)	0/3	(0.0)	1/380	(0.3)
Total	2/415	(0.5)	1/5	(20.0)	1/410	(0.2)
Prisoners						
Males	5/376	(1.3)	0/81	(0.0)	5/295	(1.7)
Females	0/34	(0.0)	0/4	(0.0)	0/30	(0.0)
Total	5/410	(1.2)	0/85	(0.0)	5/325	(1.5)
Homosexual/bisexual men	27/233	(11.6)	1/5	(20.0)	26/228	(11.4)
IVDA at the CIJ-clinic	0/11	(0.0)	0/11	(0.0)	0/0	(0.0)
Total	34/1069	(3.2)	2/106	(1.9)	32/963	(3.3)

*101 reported IV drug abuse since 1978

TABLE 2—Behavior Patterns among Homosexual/Bisexual Men by HIV-1 Antibody Status, Tijuana HIV-1 Seroprevalence Study, June–October 1988

	Seropositive		Seronegative			
Behavior	N	(%)	N	(%)	OR*	95% CI**
History of drug abuse	16/27	(59.3)	87/206	(42.3)	2.0	0.82, 4.9
History of sexually transmitted disease	15/27	(55.6)	83/206	(40.3)	1.9	0.77, 4.5
History of receptive anal sex	21/27	(77.8)	113/206	(54.9)	2.9	1.0, 8.4
History of prostitution	11/27	(40.7)	67/206	(35.5)	1.4	0.58, 3.5
History of sex in the US	12/27	(44.4)	60/206	(29.1)	2.0	0.80, 4.7
Usually or always using condoms	5/27	(18.5)	55/206	(26.7)	0.62	0.20, 1.9

*Odds ratio
**Comfield's 95% confidence limits

Within the penitentiary, 30 percent (117/410) of the prisoners reported current use of illegal drugs, 47 percent of the current users reported using IV drugs. Regular use of condoms was reported by only 4.2 percent (17/403) of the sexually active prisoners.

Discussion

This study is the largest survey reported of HIV-1 seroprevalence among potential high-risk populations in Tijuana. Nevertheless, it was impossible to reach and make precise estimates of the number and location of the homosexual/bisexual men; self-selection may have biased the seroprevalence estimated in this group. The main reason of failure to find the drug abusers selected from the clinic, was that most of them were residents of the neighboring state of California, who came to Tijuana for treatment. The almost complete lack of refusal among the prisoners and prostitutes groups actually eliminates concern of selection bias. However, bias could have been introduced when obtaining information by questionnaire regarding sensitive issues, such as sexual behavior and drug abuse.

Plausible hypotheses can be offered to explain the relatively low seroprevalence of HIV-1 infection found in this study. First, the number of partners claimed by the homosexual group is lower than that reported in high incidence places.[10–12] Second, the availability in pharmacies of sterile syringes without the requirement of a medical prescription may result in less sharing of used needles among drug abusers. However, the 1.9 percent prevalence of HIV-1 positivity in this group was similar to that estimated in various studies conducted in the Southern California area, ranging from 0 percent to 7 percent.[13] Third, the low rate among prostitutes may be attributed to the increased health consciousness of this group with monthly medical contact. Fourth, the legal weekly permission for the prisoners to receive conjugal visits by spouses or prostitutes might have lessened the possibility of widespread unprotected homosexual activity among the prisoners; however, 5 percent (19/376) of the prisoners reported history of homosexual contact during the last year. The lower availability of sterile syringes in this site and the low use of condoms indicate a serious potential of spread of HIV-1 should the virus be brought into the prisoner group.

Experience has shown that reduction in HIV-1 transmission, as well as other sexually transmitted diseases, is possible when reduction of high-risk sexual practices is induced by means of preventive intervention.[14–16] Logistic support in the implementation of preventive efforts needs to be permanently offered to nongovernmental AIDS organizations by the health authorities and coordinated and complementary action by the two countries must follow. ☐

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HIV Infection in an Ethiopian Prison

Yohannes Kebede, MD, MPH, Joyce Pickering, MD, Jane C. McDonald, MD, Kay Wotton, MD, MPH, and Debrework Zewde, PhD

ABSTRACT

HIV-1 antibody was detected in 6.0 percent of prisoners in an Ethiopian prison. HIV-1 seropositivity was strongly associated with positive VDRL status (RR = 3.7) and recent admission to prison (RR = 3.5). Seropositive prisoners were more likely to have prostitute contacts. No correlation was found between seropositivity and number of sexual contacts, long-term sex partners, or dental extractions. HIV-1 infection in this area may be of recent origin and the high prevalence may reflect rates in the surrounding community. (*Am J Public Health* 1991;81:625–627)

Introduction

Infection with human immunodeficiency virus (HIV-1) in Africa has increased rapidly. By 1988, 11,753 cases had been reported from 45 countries.[1,2] In Ethiopia, 285 AIDS (acquired immunodeficiency syndrome) cases have been reported to the National AIDS Prevention and Control Unit, by January 1990,[3] the majority (70.4 percent) from Addis Ababa.

This study was conducted in the major prison in Dire Dawa to assess the rate of HIV-1 seropositivity and associated factors. Dire Dawa is on the main communication and trade route linking Addis Ababa to the Red Sea ports of Assab and Djibouti. People in the area have high mobility and contact with neighboring regions and countries. Information obtained will lead to more relevant approaches to health prevention practices in prisons in Ethiopia and to increased awareness of the extent of HIV-1 infection in this region.

Methods

Four hundred fifty consenting prisoners from Dire Dawa District prison, present in November 1988, were enrolled. An oral closed-ended questionnaire was pretested and then administered after standardized translation into Oromingna, Somaligna, and Amharic.

Characteristics including age, sex, education, marital status, and time in prison were determined. Aspects of sexual behavior were obtained during confidential interviews.

All participants had blood samples drawn for HIV-1 and VDRL status. HIV-1 testing was done in the National

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Address reprint requests to Jane C. McDonald, MD, Division of Infectious Diseases, Montreal Children's Hospital, 2300 Tupper, Montreal, Quebec, H3H 1P3. Dr. Kebede is with the Department of Community Health, Addis Ababa University, Ethiopia; Drs. Pickering, McDonald, and Wotton are affiliated with the Department of Epidemiology and Biostatistics, McGill University, and with the Department of Community Medicine, Addis Ababa University; Dr. Zewde is with the AIDS Laboratory, National Research Institute of Health, Addis Ababa, Ethiopia. Drs. Pickering, McDonald, and Wotton are fellows of the Royal College of Physicians (Canada). This manuscript, received June 19, 1990, was revised and accepted for publication January 23, 1991.

Research Institute of Health (NRIH) in Addis Ababa using the Wellcozyme competitive enzyme linked immunosorbent assay (ELISA) (Wellcozyme anti-HTLV-III, Wellcome Diagnostics, Dartford, UK). All repeatedly reactive specimens were analyzed by Western Blot (Bio-Rad Novapath Immunoblot Assay) and considered positive if bands corresponding to at least one core protein and one envelope protein were visualized.

Syphilis serology was performed using the Venereal Disease Research Lab (VDRL) test.

Associations between HIV-1 seropositivity and sociodemographic factors and risk behaviors were examined using cross-tabulation, chi-square tests, t-test, and odds ratios with their 95 percent confidence limits. Data analysis was carried out using the SPSS statistical program.

Health education was provided to all respondents at the time of the interview. Confidential counseling to all seropositive and seronegative prisoners was offered when results were known. Those with positive VDRL tests were offered Benzathine penicillin G. with all receiving treatment.

Results

During the period November 12–23, 1988, 450 prisoners were enrolled. Of these, 27 (6.0 percent) had positive HIV-1 serology; 141 (31.6 percent) had a positive VDRL titer.

Table 1 compares relevant demographic characteristics in HIV-positive and HIV-negative prisoners. Older prisoners were 2.5 times more likely to be HIV seropositive. There was no significant association between HIV-1 seropositivity and education status, gender, or marital status. A shorter stay (< three months) was significantly associated with HIV-1 seropositivity.

Table 2 compares selected sexual behaviors. Prisoners with prostitute contact before prison were 2.5 times more likely to be HIV-1 seropositive than those prisoners who denied such contact although this did not reach statistical significance. HIV seropositivity was not significantly associated with either the number of long-term sex partners or the frequency of sexual contacts before prison. None of the prisoners reported homosexuality.

Lack of circumcision, and number of injections or dental extractions did not correlate with seropositivity (Table 3).

TABLE 1—Comparison of Selected Demographic Characteristics in HIV-Positive and HIV-Negative Ethiopian Prisoners

Demographics	HIV-positive n = 27 N (%)	HIV-negative n = 423 N (%)	Relative Risk (95% Confidence Interval)
Age (years)			
≤24	8 (4.0)	218 (36.0)	
≥25*	19 (8.5)	205 (91.5)	2.4 (1.07,5.36)
Education			
Illiterate	2 (3.7)	51 (96.3)	
Literacy program	6 (7.5)	74 (92.5)	2.0 (0.42,9.48)
Formal schooling	19 (6.0)	298 (94.0)	1.6 (0.38,6.62)
Duration (years) of Residence in Dire Dawa			
≤5	14 (7.0)	181 (93.0)	1.4 (0.68,2.93)
>5	13 (5.0)	242 (95.0)	
Travel Outside Dire Dawa			
No	4 (8.7)	42 (92.3)	
Yes	23 (5.7)	381 (94.3)	0.6 (0.24,1.81)
Prison Stay (months)			
≤3**	21 (9.3)	205 (90.7)	3.5 (1.43,8.43)
>3	6 (2.7)	218 (97.3)	

*significant variable ($p < 0.05$)
**significant variable ($p < 0.01$)

TABLE 2—Comparison of Selected Sexual Behaviors in HIV-Positive and HIV-Negative Ethiopian Prisoners

Behavior	HIV-positive n (%)	HIV-negative n (%)	Relative Risk (95% Confidence Interval)
Long-term sex partners before prison			
0	9 (6.5)	129 (93.5)	
1	11 (5.8)	177 (94.2)	0.9 (0.38,2.11)
≥2	7 (5.6)	117 (94.4)	0.9 (0.33,2.26)
Sexual contacts per month before prison			
≤1	2 (2.3)	84 (97.7)	
≥2	25 (6.9)	339 (93.1)	2.95 (0.71,12.23)
Prostitute contacts per month before prison			
0	5 (3.1)	156 (96.9)	
≥1	22 (7.6)	267 (92.4)	2.5 (0.95,6.35)

TABLE 3—Number of Injections, Dental Extractions, Immunizations, by HIV Status of Ethiopian Prisoners

	HIV-positive n (%)	HIV-negative n (%)	Relative Risk (95% Confidence Interval)
Injection during past 5 years			
0	9 (5.2)	163 (94.8)	
≥1	18 (6.5)	260 (93.5)	1.2 (0.57,2.69)
Dental extractions in past			
0	22 (6.5)	319 (93.5)	
≥1	5 (4.6)	104 (95.4)	0.7 (0.28,1.83)
Immunizations in past			
No	12 (6.0)	188 (94.0)	
Yes	15 (6.0)	236 (94.0)	1.0 (0.48,2.08)

TABLE 4—VDRL Status and HIV Infection

	HIV-positive n (%)	HIV-negative n (%)	Relative Risk (95% Confidence Interval)
VDRL Status			
Negative	10 (3.2)	299 (96.8)	
Positive*	17 (12.0)	124 (88.0)	3.7 (1.75,7.93)

*significant variable ($p < 0.001$)

The association between positive VDRL status and HIV-1 positivity was significant (Table 4).

To determine whether age was a confounder in the association between HIV positivity and other variables, chi square analyses were performed. The only significant association with age was time in prison such that young prisoners were more likely to have short stays. Because both older age and shorter stay in prison were associated with higher seropositivity rates, the effect of age on the bivariate analysis is to underestimate the shorter prison stay and higher seropositivity association.

Discussion

The seropositive prisoners in this study represent a prevalence of 6.0 percent. Among high-risk groups such as Ethiopian prostitutes, rates from 16 percent to 60 percent have been reported.[4]

Higher rates of HIV infection have been shown in prisoners than the general population. Seropositivity rates of 11 percent, 16.8 percent, and 26 percent have been reported from Switzerland, Italy, and Spain, respectively.[5] The spread of HIV infection in closed institutions is a major public health concern. An estimated 21,000 to 42,000 prisoners were infected with HIV in American State prisons in 1986 with 420 diagnosed cases of AIDS.[6-8] Although no studies have looked at the problem in African prisons, a similar situation is considered likely to exist.

The fact that short duration of stay was associated with seropositivity suggests that HIV was recently introduced into the town. Dire Dawa is on the main truck and trade route between Addis Ababa and the Red Sea, and infection could have been introduced from either direction.

Positive VDRL status was strongly associated with HIV positivity, as has been reported from Nairobi.[11] Men with HIV infection may be more susceptible to syphilis, or syphilis could change the susceptibility of the host by providing a portal of entry for the virus through a genital ulcer. Positive VDRL status may indicate greater exposure to sexually transmitted diseases including HIV[9,10] Genital ulcers, either by history or by examination, were significantly associated with HIV seropositivity in Nairobi.[11] Studies in homosexual populations have confirmed the association of genital ulcers and HIV transmission.[12]

Frequent use of prostitutes has been shown to be associated with HIV seropositivity.[11,13] In our study, a trend was apparent but not significant, possibly because of the small number of seropositive prisoners. Nevertheless, the trend, is suggestive.

A history of injections was not associated with seropositivity in this study. Results from studies in Zaire have been equivocal.[14,15] However, the continuing use of unsterilized needles in many African countries may potentially facilitate HIV transmission.

Immunizations and dental extractions were not associated with HIV seropositivity although the number of positive respondents in this area was small.

In summary, a prevalence rate of 6.0 percent was found in this prison population. The general HIV seroprevalence rate in Ethiopia is approximately 2 percent seropositivity in blood donors.[3] HIV infection appears to have been recently introduced into the area and this high rate is evidence of rapid spread in this group. Seropositivity was significantly associated with positive VDRL status and seropositive men reported more frequent prostitute contacts.

Significant educational efforts are needed to prevent further spread of HIV infection into the surrounding community. ☐

Acknowledgments

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Comparison of AIDS Knowledge, Attitudes, and Behaviors among Incarcerated Adolescents and a Public School Sample in San Francisco

Ralph J. DiClemente, PhD, Mark M. Lanier, MS, Patricia F. Horan, PhD, and Mark Lodico, PhD

ABSTRACT

Data collected from incarcerated youth (n = 113) and a public school sample (n = 802) demonstrate that both adolescent groups have a high level of AIDS knowledge. Incarcerated youth are less aware of HIV risk-reduction behaviors and report markedly higher rates of HIV risk behaviors. Incarcerated youth are at substantially increased risk of HIV infection relative to their school-based counterparts and should be a primary target of HIV prevention programs. (*Am J Public Health* 1991; 81:628–630)

Introduction

Although studies have identified the prevalence of risk behaviors among incarcerated adolescents,[1-3] as well as their knowledge about acquired immunodeficiency syndrome (AIDS),[4] at present, there are no data available comparing knowledge, attitudes, and risk behaviors between this population and a referent adolescent population. In making such a comparison, it is important to control for geographic site given recent findings indicating geographic variation in adolescents' knowledge and attitudes about AIDS,[5] prevalence of HIV (human immunodeficiency virus)-related risk behaviors,[6,7] and HIV prevalence.[8]

The present study provides a comparison of knowledge, attitudes, and HIV-related risk behaviors between adolescents in a detention facility and those in public schools in San Francisco.

Methods

In February 1988, all students enrolled in family life education classes, a required course, in nine primary high schools in the San Francisco Unified School District were asked to participate in the Centers for Disease Control (CDC) surveillance survey of adolescents' knowledge, attitudes, and risk behaviors.[9] Students completed an anonymous self-report survey during a single class period. It provided demographic information and asked about knowledge of AIDS transmission and prevention, sexual behavior, and intravenous drug use. Adolescents incarcerated at an adolescent detention facility in San Francisco also completed the survey, following the same administration protocol used with the school-based sample.

The age, ethnicity, and gender distributions of the school and incarcerated samples, while representative, demonstrate substantial differences particularly with respect to ethnicity and gender. To control for potential differences attributable to demographic variables when comparing survey responses between the samples, multivariate logistic regression analysis[10] is used to adjust for age, gender, and race.

Results

Response rates for school and incarcerated adolescents were high, 96.7 percent and 91.1 percent, respectively. Demographic characteristics of the samples are described in Table 1.

School and incarcerated adolescents generally demonstrate a high level of AIDS knowledge, with few substantial differences (Table 2). Most of the incarcerated youth correctly identify sharing of intravenous drug needles (86 percent) and sexual intercourse (90 percent) as routes of HIV transmission compared with 89 and 90 percent, respectively, for their school-based peers.

Knowledge of risk-reduction strategies identified informational gaps among both school and incarcerated youth; although incarcerated youth were less knowledgeable. While 85 percent of school youth correctly identified condom use as a means of reducing the risk of HIV transmission, only three-fourths of the incarcerated youth did so. Further, only 62 percent and 56 percent of incarcerated youth recognized that, respectively, sexual abstinence and not having sexual partners who use IV (intravenous) drugs reduce the risk of HIV infection. Among the

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Address reprint requests to Ralph J. DiClemente, PhD, Center for AIDS Prevention Studies, University of California, 74 New Montgomery Street, Suite 600, San Francisco, CA 94105 and Bayview-Hunter's Point Foundation, Multicultural Inquiry and Research on AIDS (MIRA). Dr. Lodico is also with MIRA; Mr. Lanier is with the School of Criminal Justice, Michigan State University, East Lansing, MI; Dr. Horan is with the Department of Psychology, North Carolina State University, Raleigh, NC. This paper, submitted to the Journal May 8, 1990, was revised and accepted for publication January 23, 1991.

TABLE 1—Demographic Comparison of an Incarcerated and School Sample

	Detention (n = 113)	School (n = 802)	Difference (95% CL)
Gender			
Male	76.4	48.5	27.9 (19.4, 36.4)
Female	23.6	51.5	27.9 (19.4, 36.4)
Age (years)			
≤14	16.1	10.6	5.5 (0, 12.6)
15–16	53.6	74.3	20.7 (11.0, 30.4)
≥17	30.4	15.2	15.2 (6.4, 24.0)
Race			
White	11.0	11.4	0.4 (0, 6.6)
Black	65.1	12.4	52.7 (43.6, 61.8)
Hispanic	10.1	13.7	3.6 (0, 9.6)
Asian	5.5	55.9	50.4 (45.0, 55.8)
Other	8.3	6.6	1.7 (0, 7.1)

95% CL = 95% Confidence Limits

school-based sample, correct responses were 80 percent and 72 percent, respectively.

Attitudes about AIDS

Both samples strongly endorsed the need for AIDS education (Table 2). Marked differences were identified with respect to perceived risk of HIV infection, as 69.4 percent of incarcerated youth compared with 45.2 percent of the school-based sample perceived themselves susceptible to infection. Awareness of AIDS information resources or where to receive an antibody test was low, although incarcerated youth are more knowledgeable than their school-based peers.

Prevalence of Risk-Taking Behavior

Substantial differences were identified between the samples with respect to prevalence of HIV risk behaviors (Table 3). Almost all incarcerated youth report being sexually experienced compared with approximately 28 percent of the school sample. Most incarcerated youth (73 percent) report two or more sex partners within the past year compared with approximately 8 percent of the school sample. A large proportion of incarcerated youths initiate sex at an early age as approximately 52 and 26 percent of the incarcerated and school samples, respectively, report sexual onset at 12 years of age or younger. Condom use among both samples is inconsistent, although incarcerated youth report being less likely to be consistent condom users. Intravenous drug abuse is prevalent, especially among incarcerated youth, although roughly sim-

TABLE 2—Knowledge about AIDS and Risk-Reduction Strategies

	% Correct		Odds Ratio[a]	95% Confidence Limits
	Detention	School		
Statement				
—You can protect yourself from being infected with the AIDS virus.	80.7	84.5	.56	.31, 1.0
—A person can get AIDS from sharing needles used to inject drugs.	85.6	88.6	.67	.35, 1.3
—You can tell if a person has the AIDS virus by looking at the person.	64.2	70.6	.75	.48, 1.2
—Any person infected with the AIDS virus can infect someone else during sexual intercourse.	89.9	93.5	.56	.26, 1.2
—A pregnant woman who has the AIDS virus can infect her unborn baby.	90.8	89.1	1.3	.63, 2.8
—A person can get AIDS by having sexual intercourse.	89.9	89.8	.88	.42, 1.8
—There is a cure for AIDS.	72.5	76.0	.57	.34, .96
—Only gay men can get AIDS.	88.1	92.2	.59	.29, 1.2
—Blood transfusions are now generally safe.	25.7	31.5	.59	.36, .97
Risk-Reduction Strategies				
People can reduce their chances of becoming infected with the AIDS virus by:				
—not having sexual intercourse	62.4	80.3	.38	.24, .61
—using condoms during sexual intercourse	75.9	84.5	.43	.25, .74
—urinating after sexual intercourse	50.9	53.9	.65	.42, 1.0
—having sexual intercourse with one person not infected with AIDS virus	61.1	69.0	.61	.39, .96
—not having sexual intercourse with a person who uses illegal drugs that can be injected	56.0	71.6	.44	.28, .70
—taking birth control pills	74.3	73.7	.80	.49, 1.3
Attitudes about AIDS				
Should students be taught about AIDS in school?	88.4	89.0	1.05	.50, 2.2
Should a student with AIDS have the right to go to school?	46.4	57.0	.64	.44, .99
Would you be willing to go to class with a student with AIDS?	42.3	48.7	.72	.47, 1.1
Do you think you can get AIDS?	69.4	45.2	2.25	1.4, 3.6
Can you keep from getting AIDS?	81.3	77.3	.91	.51, 1.6
Do you know where to get correct information about AIDS?	60.7	53.2	1.01	.65, 1.6
Do you know where to get tested for the AIDS virus?	54.5	34.7	1.81	1.2, 2.8

[a] Odds ratio adjusted by age, gender, and race.

TABLE 3—Prevalence of Sexual and Drug-related Risk Behaviors

Behavior	% Detention	% School	Odds Ratio[a]	95% Confidence Limits
Proportion sexually active	99.1	28.3	5.5	3.3, 9.4
Number of sex partners				
≥3 Lifetime	84.4	15.1	7.4	4.4, 12.4
≥2 Past year	72.9	8.1	9.4	5.6, 15.7
Age at sexual onset				
≤12 years of age	52.1	25.9	2.5	1.5, 4.4
Proportion reporting:				
Consistent condom use[b]	28.6	37.3	0.6	.35, 1.1
Use intravenous drugs[c]	12.9	3.7	3.3	1.5, 7.4

[a] odds ratio adjusted by age, gender, and race.

[b] among sexually active, defined as "always" using condoms.

[c] proportion reporting lifetime prevalence of intravenous drug use.

ilar proportions of intravenous drug users in each sample report sharing needles.

Discussion

Incarcerated adolescents compared with a school sample demonstrate similar levels of AIDS knowledge, although they are less likely to be knowledgeable of prevention strategies—especially the protective value of condoms. Incarcerated youth were more likely to be sexually active, they initiated sex at an earlier age, report more sexual partners, and a smaller proportion use condoms consistently. This combination of behaviors will greatly increase the risk of HIV infection in this population. Likewise, the higher prevalence of intravenous drug abuse among incarcerated youth will also increase their risk.

The findings indicate an urgent need to implement effective AIDS prevention programs for these high-risk youth. AIDS prevention programs specifically targeted for incarcerated youth will have to overcome particular barriers to behavior change, such as low motivation for self-protection and lack of self-esteem.[11] Without prompt intervention, the potential for rapid spread of HIV infection in this population and transmission to other adolescents and adult populations appears substantial. □

Acknowledgments

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The Donation and Sale of Blood by Intravenous Drug Users

Dale D. Chitwood, PhD, J. Bryan Page, PhD, Mary Comerford, MSPH, James A. Inciardi, PhD, Clyde B. McCoy, PhD, Edward Trapido, ScD, and Shenghan Lai, PhD, MPH

ABSTRACT

In spite of efforts to dissuade intravenous drug users (IVDUs) from donating or selling blood, some continue to do so. As part of a longitudinal study, 915 IVDUs in South Florida were interviewed concerning their history of donation or sale of blood and tested for antibodies to HIV-1 and HTLV-I/II. Approximately 17 percent had either donated or sold blood during 1985 through 1988; most contributors (80.4 percent) sold to commercial blood services. IVDUs who had donated/sold blood were more likely to be male and not in drug treatment than were those who had not contributed blood. IVDUs not in treatment at the time of interview were more likely than IVDUs in treatment to have sold blood. Of those who had donated/sold blood since 1985, 19.6 percent subsequently tested positive for antibodies to HIV-1 and 5.7 percent were positive for antibodies to HTLV-I/II. Increased effort is required to screen prospective donors and sellers, particularly at commercial blood banks. (*Am J Public Health* 1991;81:631–633)

Introduction

Because HIV infectivity precedes antibody formation,[1,2] blood service organizations must rely on self-reported history and physical signs to exclude intravenous drug users (IVDUs) from donor pools. The exclusion of IVDUs is also important because they are at high risk for the human lymphotropic viruses (HTLV-I/II),[3,4] which can be transmitted through infected blood[5,6] but for which blood services do not routinely screen.

In spite of increased donor exclusion efforts, IVDUs continue to donate and sell blood.[7] Nelson and colleagues[8] studied the blood donation histories of a predominately Black cohort of IVDUs from the Baltimore metropolitan area and reported that 27.1 percent had donated or sold blood between 1978 and 1988. Data from other geographic areas have not been reported.

We studied a tri-ethnic cohort of IVDUs in Miami-Dade County, Florida, which has the highest number of cases of AIDS (acquired immunodeficiency syndrome) attributable to parenteral drug use in the Southeastern United States.[9] As part of this investigation, we assessed the seroprevalence of HIV-1 and HTLV-I/II among the study sample, compared IVDUs who donated/sold blood with those who did not, compared donors with sellers, and examined the association of drug treatment with the donation/sale of blood.

Methods

Between June 1987 and August 1988, 915 IVDUs in South Florida were enrolled in a longitudinal study to determine HIV-1 serostatus and associated risk behaviors (22.8 percent were HIV seropositive at baseline). Seven hundred and four participants were recruited from drug treatment facilities in South Florida (eight methadone maintenance clinics, three residential drug-free programs, and one residential detoxification facility). An additional 211 subjects not enrolled in a drug treatment program were recruited directly from the streets. All subjects had used a drug or drugs intravenously at some time since January 1, 1978. All participated on a voluntary basis, gave informed consent, were tested for HIV-1 serostatus, and received pre- and post-test counseling. As part of a structured questionnaire, participants answered questions concerning their most recent donation or sale of blood. All were assured these data would be kept fully confidential and would not be disclosed to anyone including treatment center staff. Participants were paid a fee after completion of venipuncture and interview.

All sera were tested for the presence of antibodies to HIV-1. Residual sera, available on 874 of the study participants, were tested subsequently for antibodies to HTLV-I/II. Because of cross-reactivity, the assay did not differentiate between antibodies to HTLV-I or HTLV-II.

Serum was determined to be positive for antibodies to HIV-1 after two positive ELISAs (enzyme-linked immunosorbent assay) determined with Abbott Enzyme Immunoassay and a confirmatory positive Western Blot following the Centers for Disease Control protocols using viral lysate from Hillcrest Biologicals. A radioimmune assay utilizing an in-house ^{125}I-ionated core protein (p24), purified from human virus, was the antigen for detecting antibody to HTLV-I/II. Serum samples were tested at dilutions of 1:10 and 1:100, which bracket the dilution levels of maximum sensitivity for sera of all titers. Sera were evaluated in terms of percent-bound counts using a high-titered reference se-

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From the Comprehensive Drug Research Center, Department of Oncology, Biopsychosocial Center for the Study of AIDS, and the Department of Psychiatry, University of Miami School of Medicine, all authors except Inciardi who is with the Division of Criminal Justice, University of Delaware, Newark, DE. Address reprint requests to Dale D. Chitwood, PhD, Research Associate Professor, School of Medicine, University of Miami, 1550 N.W. 10th Avenue, Room 301, Miami, FL 33136. This paper, submitted to the Journal May 9, 1990, was revised and accepted for publication November 20, 1990.

Editor's Note: See also related editorial on page 561 of this issue.

TABLE 1—HIV-1 and HTLV-I/II Seroprevalence and History of Donation/Sale of Blood among IVDUs 1985–1988

	HIV-1*		
	+	−	Prevalence (%)
Never Gave	178	583	23.4
Donated	4	26	13.3
Sold	26	97	21.1
	HTLV-I/II**		
	+	−	Prevalence (%)
Never Gave	74	659	10.1
Donated	1	24	4.0
Sold	7	109	6.0

Four individuals who had sold or donated blood were seropositive for antibodies to both HIV-1 and HTLV-I/II.

*One subject tested indeterminate and was excluded from the analysis.

**Residual blood was not available for HTLV-I/II testing for 41 study participants.

TABLE 2—Multiple Logistic Regression Analysis of Behavior of Selling and/or Donating Blood

	Odds Ratio	95% CI
Age (years)		
<30	1.00	
>30	0.79	(0.55, 1.13)
Gender		
Male	1.00	
Female	0.26	(0.19, 0.36)
Ethnicity		
Black		
No	1.00	
Yes	0.96	(0.65, 1.41)
Hispanic		
No	1.00	
Yes	0.63	(0.38, 1.03)
Treatment Status		
Not in treatment	1.00	
In treatment	0.59	(0.39, 0.90)

rum at 1:100 dilution as the 100 percent standard. Sera were scored as positive if the percent-bound counts were greater than 5 percent at the 1:100 dilution.

Logistic regression analysis was used to evaluate the possible association between demographic characteristics of IVDUs and their behavior regarding the donation or sale of blood since 1985.

Results

Of the 915 IVDUs interviewed, 153 (16.7 percent) reported that they had donated or sold blood since January 1, 1985; four out of five (80.4 percent) of these 153 IVDUs had sold blood. A total of 22.8 percent of the study population tested positive for HIV-1 (Table 1). Of the 153 IVDUs who had contributed blood, 19.6 percent tested positive for antibodies to HIV-1 at the time of interview.

Of the subjects who had donated or sold blood, 5.7 percent of those tested for HTLV-I/II were positive. Serostatus at the time they donated/sold their blood is unknown.

When IVDUs who had donated or sold blood were compared with IVDUs who had not done so in a logistic regression, males and IVDUs who were not in treatment were more likely to have donated/sold blood (Table 2). No differences were observed by ethnic or age group. The demographic characteristics of IVDUs who had donated and those who had sold blood were also compared (data not shown). Donors were differentiated from sellers only by treatment status. Of those in treatment who had sold/donated blood, 68.5 percent had sold; 98.8 percent of the IVDUs from the street had sold rather than donated.

Discussion

Although participation in the study was voluntary, results indicate that the donation and sale of blood by IVDUs is not restricted to the Northeastern United States, and that the sale of plasma or blood accounts for almost all IVDU involvement in the blood supply. The percentage of IVDUs in South Florida (16.7 percent) who donated or sold blood between 1985 and 1988 was virtually identical to the 15.0 percent reported by Nelson and colleagues[8] for a cohort of IVDUs in Baltimore. At least four out of five IVDUs from both Miami (80.4 percent) and Baltimore (88.1 percent) who contributed blood during that time period had sold blood or plasma at a commercial service rather than donated at a voluntary blood center.

Blood services utilize blood screening procedures to ensure that the supply is as safe as possible, but the elimination of high-risk donors remains an unfinished task in the overall plan to protect the blood supply. Since the initiation of blood screening for HIV antibodies and the utilization of the Cohn fractionation process and pasteurization procedures, the risk of a recipient receiving HIV infected blood or blood products has diminished significantly, but has not disappeared.[10–12] Screening for antibodies to HTLV-I/II is not routinely conducted. Several IVDUs who donated or sold blood currently test positive for HIV-1 and/or HTLV-I/II. It must be emphasized that there is no way to ascertain the serostatus of these IVDUs at the time of their last donation or sale. Whether any of the blood contributed by these IVDUs became part of the blood supply is unknown, but it is evident that members of this high-risk group continue to donate/sell blood.

The elimination of the sale of blood by IVDUs is a difficult matter. As long as commercial centers exist, IVDUs will attempt to sell their blood. Two characteristics of commercial blood centers facilitate this selling behavior: commercial blood centers which advertise that they pay cash for blood or plasma attract active drug users who do not have stable sources of income; commercial blood centers often are located in inner-city neighborhoods where poverty and drug use are prevalent.

Commercial blood centers must take a more careful approach in the selection of potential sellers. It is imperative that the screening process at commercial blood services be intensified to identify potential sellers who inject drugs or have a history of IV drug use. Commercial blood centers should employ highly qualified individuals, who have had experience with IVDUs, to screen potential sellers for eligibility. Included in the screening process should be a rigorous check for track marks (scarring caused by repeated injections) on arms and other indications of IV drug use.

Drug treatment programs can play an important role in the reduction in the number of IVDUs who donate or sell blood. Educational intervention with past sellers and donors in treatment could have an immediate positive impact.

Blood services, both voluntary and commercial, have an enormous responsibility to exclude IVDUs from their contributor pool. It is incumbent upon these services to staff the blood centers with highly qualified persons who can identify and exclude high-risk persons who continue to donate or sell blood. ☐

Acknowledgments

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Hyperendemic Urban Blastomycosis

A. Cesare Manetti, MD, MPH

ABSTRACT

Within the United States, blastomycosis is primarily endemic in the Mississippi and Ohio River valleys, but reliable figures as to its prevalence and incidence are scarce. A review of pathology department files and medical records from three hospitals in Rockford, Illinois, revealed 32 documented cases of blastomycosis in an eight and one-half year period beginning in March 1981. Twenty-three of these cases were urban since they occurred within the Rockford census tract boundaries—an incidence more than 3.5 times greater than expected in an endemic area. Addresses of patients in 18 of the 23 cases were concentrated at the northeast and southwest ends of the city. Implications are discussed. (*Am J Public Health* 1991;81:633–636)

Introduction

North American Blastomycosis is an elusive, cosmopolitan, systemic infection caused by the dimorphic fungus *Blastomyces dermatitidis.* First described by Gilcrest[1] in 1894, the infection became known as "the Chicago disease" because of the number of cases seen from that area. Later it was referred to as North American blastomycosis as the extent of its wider distribution in the United States and Canada was discovered. Today the disease is correctly and simply referred to as blastomycosis, and its etiologic agent is known to be endemic in Europe, the Middle East, Africa and India. In the Mississippi and Ohio River valleys, where it is primarily endemic, blastomycosis occurs sporadically in humans, dogs and other mammals. Blastomycosis is considered rare, but its true incidence is not known since it is not reportable in most states.

Outbreaks of blastomycosis are seldom observed, and those reported have offered unique opportunities to investigate the source of the infection. The nine outbreaks that have been described to date have mostly been associated with outdoor activities.[2,3] All but one have occurred in non-urban locations, and six were observed in the Great Lakes region. In three of the outbreaks, it was possible to isolate *B. dermatitidis* from rotten wood at exposure sites, mostly in an environment around waterways.[3,4]

One of the most baffling aspects of blastomycosis has been the difficulty encountered in demonstrating where the oganism lives in nature and where humans and animals contract the infection. Any unusual concentration of blastomycosis cases is of interest since it may offer some clues as to the natural habitat of *B. dermatitidis.* This is a report of two hyperendemic foci of blastomycosis in an urban setting.

Methods

Blastomycosis is not a reportable illness in the State of Illinois, nevertheless, a sufficient number of cases came to the attention of the director of the Winnebago County Health Department to suggest the need for an investigation of a possible underlying problem. A search of the files of three Rockford hospital pathology departments and of their respective medical records uncovered 32 cases of blastomycosis. They had all been diagnosed by one or more of three modalities, namely culture, histologic sections and cytologic examina-

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From the Department of Community Health Sciences, School of Public Health, University of Chicago; and the Department of Pathology, University of Illinois College of Medicine at Rockford. Address reprint requests to A. Cesare Manetti, MD, MPH, P.O. Box 4508, Rockford, IL 61110-4508. This paper, submitted to the Journal May 15, 1990, was revised and accepted for publication November 2, 1990.

TABLE 1—Rockford Hospitals: Blastomycosis Cases [March 1981–August 1989]

Case No.	Specimen (Date/Type)	Census Tract	Age/Race/Sex Occupation	Chief Complaint	Signs, Symptoms and Duration	Initial diagnosis/ Remarks
1	3/81 lung biopsy	2	39, Wh, M factory clerk	Cough	Weight loss, fever	Unresolving pneumonia/ thoracotomy
2	3/81 sputum cytology	28	45, Bl, M factory worker	Cough	Weight loss 5 weeks	Unresolving pneumonia
3	4/81 lung biopsy	13	24, Wh, F secretary	Hemoptysis	Pleural pain	Unresolving pneumonia
4	7/82 skin biopsy	36.01	20, Wh, F	Skin problem		Dermatitis
5	12/82 ear tissue	out of Rockford	25, Wh, M salesman	Ear pain, debility	Weight loss, 3 months	Mastoiditis/ mastoidectomy
6	2/83 bronchial biopsy	out of Rockford	31, Wh, F	Chest pain	Weight loss,	Unresolving pneumonia/ bronchoscopy
7	3/83 bronchial lavage	1.05	31, Wh, M grinder	Cough	Hemoptysis	Unresolving pneumonia
8	3/83 sputum, abscesses	1.05	44, Wh, M teacher	Cough	Weight loss, fever, abscesses lower extremity	Unresolving pneumonia
9	7/83 sputum	out of Rockford	31, Wh*, M factory worker	Cough	Weight loss, fever, chest pain	Unresolving pneumonia
10	12/83 tissue	2	42, Wh, M stock rm clerk	Hemoptysis	Weight loss 4 weeks	Unresolving pneumonia/lobectomy
11	2/85 lung biopsy	out of Rockford	55, Wh, F laborer	Chest pain	7 months	Unresolving pneumonia
12	9/85 lung aspirate	2	33, Wh, M machinist	Cough	Fatigue, chest pain, weight loss	Unresolving pneumonia/suspect blastomycosis
13	4/86 bronchial aspirate	37.03	48, Wh, M mechanic	Cough	Weight loss 2 months	Unresolving pneumonia
14	6/86 lung tissue	1.03	43, Wh, M	Cough	Chills, fever 2 months	Unresolving pneumonia
15	7/86 sputum	out of Rockford	55, Wh, M farmer	Cough	Weight loss, fever	Unresolving pneumonia
16	7/86 lung biopsy	out of Rockford	50, Wh, M factory worker	Cough	Flank pain weeks	Unresolving pneumonia/lung cancer
17	8/86 skin, bone and sputum	out of Rockford	44, Wh, M school coach	Wrist pain	Chest problems months	Synovitis, dermatitis, osteomyelitis, pneumonia
18	8/86 sputum	24	39, Bl, M factory worker	Cough	Chest pain, fever 4 weeks	Unresolving pneumonia/ bronchoscopy
19	4/87 sputum, skin	2	41, Wh, F unemployed	Depression	Chest pain, wt loss, fever 4 months	Psychiatric care, lung infiltrates, skin lesions: dermatitis
20	5/87 sputum	27	47, Bl, M furnace repair	Hemoptysis	Cough, skin lesions	Lung mass, dermatitis
21	10/87 sputum	4.01	89, Wh, M factory worker (retired)	Chest pain		Lung cavity by X-ray
22	11/88 sputum	1.04	38, Wh, F	Hemoptysis	Cough, chills	Unresolving pneumonia
23	12/88 pus, sputum	26	33, Bl, M maintenance worker	Calf abscess	Chills, fever, hemoptysis, weight loss	Abscess pneumonia
24	12/88 sputum	24	26, Bl, M	Cough		Pneumonia
25	1/89 skin biopsy	20	53, Wh, F	Skin lesions	multiple sites	Expired; postmortem miliary blastomycosis
26	2/89 sputum	32	27, Bl, F factory worker	Side pain	Wt loss, fever 5 weeks	Unresolving pneumonia
27	3/89 sputum	out of Rockford	59, Wh, M groundskeeper	Cough	bronchitis, 8 weeks	
28	3/89 sputum	out of Rockford	36, Wh*, F	Cough	Fever, chills	Unresolving pneumonia
29	4/89 joint fluid	34	27, Bl, M construction worker	Elbow pain	Swelling, fever, cough, wt loss 4 months	Osteomyelitis, pneumonia
30	5/89 sputum	32	44, Bl, M unemployed	Cough	Weight loss 6 weeks	Pneumonia
31	7/89 bronchial lavage	5.06	51, Wh, F	Chest pain	Cough, hemoptysis, weight loss	Unresolving pneumonia
32	7/89 trans-bronchial biopsy	22	29, Bl, M	Cough	Chills, fever, weight loss 4 weeks	Unresolving pneumonia

*Hispanic surname

Listing of 32 Rockford blastomycosis cases showing time of diagnosis, census tract and, where applicable, age, race, sex, occupation, chief complaint, signs and symptoms, duration of symptoms at time of initial diagnosis and initial working diagnosis.

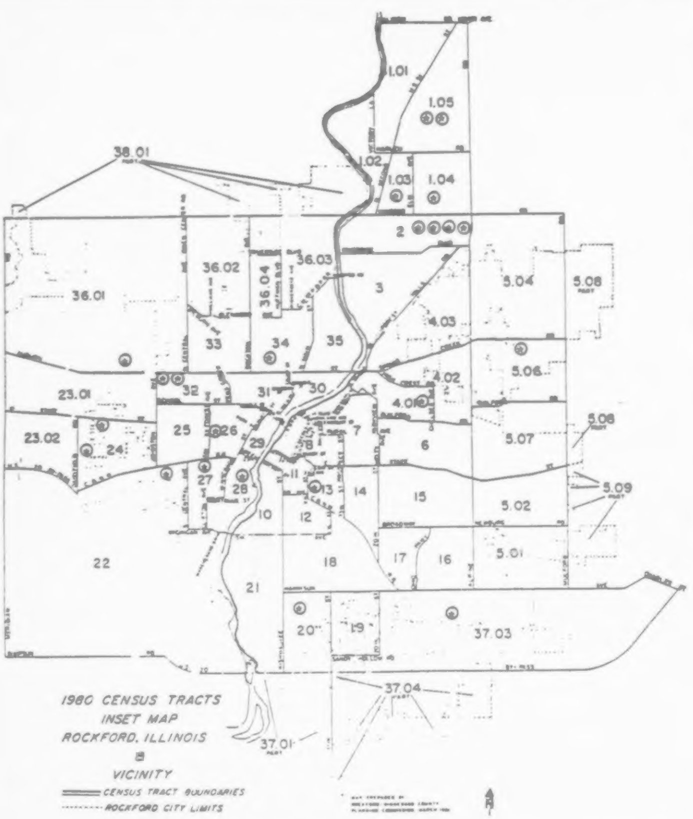

FIGURE 1—1980 Census Tract Inset Map for Rockford and Vicinity in which the Residences of the 23 Urban Rockford Blastomycosis Cases Are Shown Highlighted by a Circled Asterisk.

TABLE 2—Blastomycosis in the City of Rockford (Southwest side) 1980 Population

Census Tract	Total	(Black)	Year	Race	Sex
36.01	2,267	(38)	1982	W	F
34	4,944	(45)	1989	B	M
32	3,965	(1,217)	1989	B	M
			1989	B	F
26	4,452	(2,501)	1988	B	M
24	3,331	(2,402)	1986	B	M
			1988	B	M
22	5,227	(1,960)	1989	B	M
27	3,795	(2,108)	1989	B	M
28	2,153	(1,310)	1981	B	M
Total	30,134	(11,581)		M/F = 8/2	

The 10 cases from one of the two high concentration areas of town are shown according to their specific location, year of occurrence, race and sex. Total and Black population figures per census tract are also given.

tion, during a time span of eight and a half years (from March 1981 to August 1989).

The cases are listed in Table 1 by the month and year when the diagnosis of blastomycosis was first established and by the type of specimen that yielded the diagnosis. The next column gives the census tract in which each Rockford patient lived at the time of initial diagnosis (the patients from other communities are listed as "out-of-Rockford"). Listed in other columns are age, race, sex, occupation, chief complaints, principal signs, symptoms and their duration at the time of initial diagnosis. The initial working diagnosis and other pertinent remarks are given in the last column.

Results

Nine of the 32 patients were referred to Rockford hospitals from outlying areas. Twenty-three patients resided within the census boundaries of Rockford; we will refer to these as the urban areas.

Figure 1 is the 1980 census tract inset map for Rockford and vicinity. In this map, the residences of the 23 urban Rockford cases are highlighted by a circled asterisk.

The distribution of the residences of the 23 Rockford patients was as follows: eight were concentrated in the northeast, 10 were concentrated on the southwest side, and the remaining five were scattered throughout the rest of the city.

The population figures quoted in this study are taken from the 1980 United States Census published in April 1981. According to the Rockford Community Development Department, the population of the city has not appreciably increased or shifted since that time.

Tables 2 and 3 show the 10 cases from the southwest side of town and the eight cases from the northeast side of town according to their census tract, year of occurrence, race and sex, as well as the total population and the black population figures per the 1980 census tract.

Based on the 1980 Rockford census population of 139,712, the annual incidence of blastomycosis in Rockford can be estimated at 1.94/100,000 (23/139,712 × 1/8.5).

Discussion

The following are yearly blastomycosis incidence rates for some states where blastomycosis is endemic[2]:

Arkansas	0.43/100,000
Kentucky	0.44/100,000
Mississippi	0.61/100,000
Wisconsin	0.48/100,000

The estimated incidence rate in Rockford taken as a whole is 3 to 4.5 times the incidence rate for endemic areas, while the estimated incidence in two parts of town is much higher. The annual blastomycosis incidence rates for the south-

TABLE 3—Blastomycosis in the City of Rockford (Northeast side) 1980 Population

Census Tract	Total	(Black)	Year	Race	Sex
1.05	4,242	(81)	1983	W	M
			1983	W	M
1.03	2,654	(5)	1986	W	M
1.04	3,409	(23)	1988	W	F
2	3,807	(8)	1981	W	M
			1983	W	M
			1985	W	M
			1987	W	F
Total	14,112	(117)		M/F = 6/2	

The eight cases from the other high concentration area of town are shown according to their location, year of occurrence, race and sex. Total and Black population figures per census tract are also given.

west side of town is estimated at 3.9/100,000 and that for the northeast area of town is 6.67/100,000—rates which are, respectively, 6 to 9 and 10 to 15 times those reported for endemic blastomycosis.

It should be noted, however, that while the northeast cases all occurred among the predominantly White population, nine out of 10 southwest side cases occurred among the Black minority population, an incidence among Blacks of 9.14/100,000. By contrast, the incidence among Whites inhabiting the same southwest census tract is 0.63/100,000, which is within a reasonable range of the expected endemic incidence.

By the same token, since only five cases of blastomycosis occurred during the same period of time in the population of Rockford living outside of the two high incidence areas, the calculated incidence for these 95,466 individuals is a similar 0.62/100,000.

The present report draws attention to the fact that there can exist foci of a persistent, high concentration of blastomycosis cases in an urban setting. Because of the privileged nature of the hospital records, we were prevented from contacting patients directly to explore further the possible causes of this unusual concentration of cases.

Blastomycosis is a serious, insidious, sometimes indolent disease that frequently goes unrecognized. Eleven of the 50 US states require mandatory reporting of blastomycosis,[5] but many states in a well known endemic area do not, as in the case of Illinois, Indiana, Iowa, Kentucky, Michigan, Missouri, New York and Tennessee.

Given the complexity and serious nature of the condition, we strongly recommend that blastomycosis be included among the reportable infections in Illinois and in neighboring states. ☐

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Use of a Medical Center's Computerized Health Care Database for Notifiable Disease Surveillance

Margaret Watkins, BSN, MPH, Sandra Lapham, MD, MPH, and Wendy Hoy, MB, BScMed

ABSTRACT

The sensitivity of a medical center's inpatient and outpatient database to detect notifiable diseases was examined. Only 53 percent of inpatient and 7 percent of outpatient laboratory-confirmed cases of shigellosis, salmonellosis, giardiasis, and hepatitis were identified by an automated search for matching diagnosis codes. Reasons for lack of sensitivity include nonavailability of laboratory results at the time of diagnosis assignment, use of a standardized encounter form with limited preselected diagnosis codes, and pre-empting of the infectious disease diagnosis by other diagnoses. (*Am J Public Health* 1991; 81:637–639)

Introduction

Most communicable disease surveillance systems in the United States are based on passive reporting by health care providers. Because these have major limitations, such as underreporting and reporting bias,[1,2] investigators continue to search for alternative data sources. Promising among these are computerized hospital discharge databases. Although hospital discharge data have been used for disease surveillance,[3,4] questions remain about their validity and sensitivity. This study examines the sensitivity of one medical center's computerized inpatient and outpatient tracking and billing system when used for notifiable disease surveillance. Also assessed was the medical center's completeness of notifiable disease reporting to the health department.

Methods

Lovelace Medical Center in Albuquerque, New Mexico is a 235-bed hospital and multispecialty group practice with eight "satellite" family practice and urgent care centers. During the study period August 1, 1986–July 31, 1987, Lovelace had approximately 10,000 inpatient admissions and 540,000 outpatient visits. A computerized billing system is used to record International Classification of Diseases, 9th Ed, Clinical Modification (ICD-9-CM)[5] diagnosis codes, as well as demographic, diagnostic and billing information. Inpatient codes are assigned shortly after discharge by Medical Records Department staff, based on providers' documentation in discharge summaries and progress notes. Up to 30 diagnosis codes are captured by the system. Outpatient visit codes (both "scheduled" and "urgent care") are assigned at the time of the visit by the medical provider, using a standardized encounter form with pre-selected diagnostic codes. A maximum of three codes is allowed on the outpatient form.

To determine the sensitivity of the ICD-9-CM coding system and to determine the reporting rates to the health department, laboratory-identified cases of five notifiable diseases were used as the "gold standard." These included shigellosis, salmonellosis, giardiasis, and hepatitis A and B. Inpatient and outpatient databases were searched for matching ICD-9-CM codes, and cases of notifiable diseases reported by Lovelace personnel to the New Mexico Health and Environment Department (NMHED) were ascertained. To assess reasons for discrepancies among these data sources, medical record reviews were performed on cases that were laboratory-identified but not reported to the NMHED (n = 44) and cases that were laboratory-identified but not ICD-identified (n = 207). Forty-two of the former (two records not found) were reviewed. Due to resource limitations, a random sample of 60 (29 percent) of the latter cases were reviewed.

Attempts were made to review all non-laboratory-identified cases that were either identified by ICD-9-CM code search (n = 9) or reported to the NMHED (n = 6).

Results

Sensitivity of ICD-9-CM Diagnosis

Each of the three data sources yielded cases that would have otherwise gone undetected (Figure 1). Only about half the inpatient and only 7 percent of the outpatient laboratory-confirmed cases were identified by the matching ICD-9-CM codes (Table 1). However, the ICD-9-CM search identified nine cases—all among outpatients—which were not laboratory-confirmed. None of these were reported to the NMHED. One was a symptomatic contact of a known case, two were

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Address reprint requests to Margaret Watkins, BSN, MPH, Research Associate, Clinical Studies Division, Lovelace Medical Foundation, 2441 Ridgecrest SE, Albuquerque, NM 87108. Dr. Lapham is an epidemiologist and Dr. Hoy is director of the Division, both at the Foundation. This paper, submitted to the Journal May 2, 1990, was revised and accepted for publication December 20, 1990.

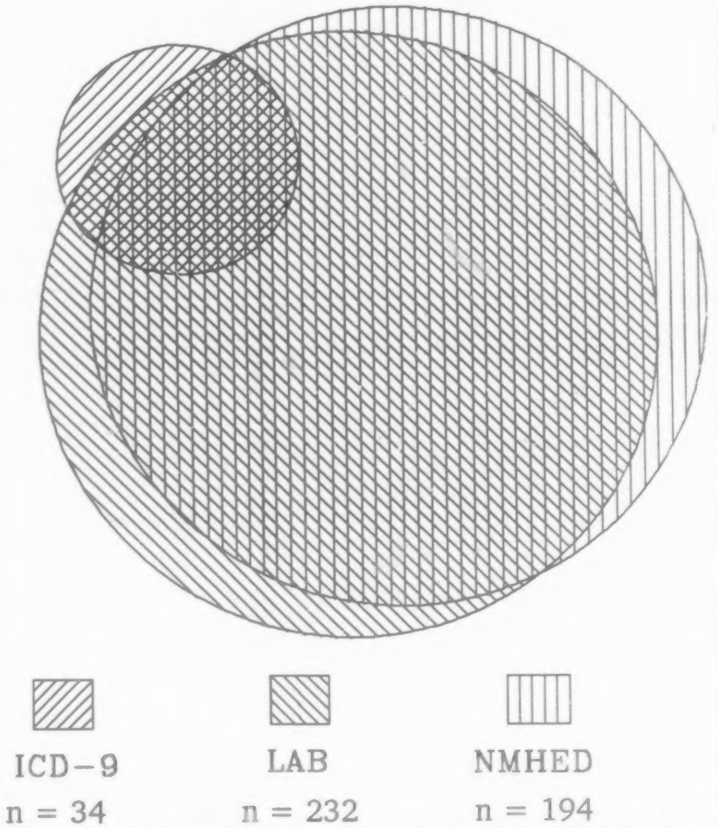

FIGURE 1—Notifiable Disease Cases[a] Identified by Laboratory Confirmation, ICD-9-CM[b] Diagnosis Search Cases Reported to the NMHED,[c] Lovelace Medical Center 8/1/86–7/31/87

a) Salmonellosis, Shigellosis, Giardiasis, Hepatitis A, and Hepatitis B.
b) ICD-9-CM—International Classification of Diseases, 9th ed, Clinical Modification.
c) NMHED—New Mexico Health and Environment Department, Santa Fe, NM.

TABLE 1—Inpatient and Outpatient Laboratory-Confirmed Diagnoses at Lovelace Medical Center: Percent Reported to NMHED[a] and Detected by ICD-9-CM[b] Codes August 1, 1986–July 31, 1987

	Reported to the NMHED[a]		ICD-9-CM[b] Search	
	Inpatient (%)	Outpatient (%)	Inpatient (%)	Outpatient (%)
Salmonella	3/3 (100)	53/54 (98)	2/3 (67)	3/54 (4)
Shigella	1/1 (100)	23/23 (100)	0/1 (0)	1/23 (4)
Giardia	2/2 (100)	25/44 (57)	1/2 (50)	3/44 (7)
Hepatitis A	4/4 (100)	39/55 (71)	3/4 (75)	3/55 (5)
Hepatitis B	9/9 (100)	29/37 (78)	4/9 (44)	6/37 (16)
TOTAL	19/19 (100)	169/213 (79)	10/19 (53)	15/213 (7)
TOTAL inpatient and outpatient	188/232 (81)		25/232 (11)	

[a] NMHED—New Mexico Health and Environmental Department, Santa Fe, NM.
[b] ICD-9-CM—International Classification of Diseases, 9th Ed, Clinical Modification.

chronic or previously diagnosed cases, two were follow-up visits, and one was a case in which the diagnosis—based solely on symptoms—was questionable. (Three records were unattainable.)

Reporting to the Health Department

All inpatient cases and most outpatient laboratory-confirmed cases were reported to the NMHED. Six cases were reported that were not laboratory confirmed; these were symptomatic contacts of known cases or cases in which the diagnosis was based solely on symptoms.

Discussion

Notifiable disease surveillance based solely on ICD-9-CM diagnosis codes was not very sensitive.

In the outpatient setting, miscoding on the initial visit is partly explained by the fact that the laboratory diagnosis has not been established or confirmed. However, medical record review revealed that at the follow-up visit, when positive laboratory results were documented, matching ICD-9-CM codes were still not recorded in approximately half of the cases. Another explanation for the failure to assign appropriate ICD-9-CM codes was the bias introduced by the use of a standardized encounter form with a limited number of preselected diagnosis codes, some of which were nonspecific. The provider seldom uses the option of writing in a diagnosis other than those given on the form. Changes in the encounter form content or usage protocols would be needed before outpatient diagnosis codes would be useful for infectious disease surveillance.

Only half of the inpatient cases were detected by matching ICD-9-CM codes. Review of all the laboratory-identified cases not identified by ICD-9-CM revealed two main reasons for this discrepancy: 1) as with outpatient cases, laboratory results were not always available at the time of discharge, and 2) competing diagnosis; i.e., the infectious disease diagnosis was pre-empted by other diagnoses in complex cases. For example, one patient readmitted with multiple complications after cardiac surgery had 13 discharge diagnoses that did not include his nosocomially-acquired hepatitis B.

Rates of reporting of laboratory-confirmed cases to the NMHED were high. Current policy not only requires that reports of laboratory-confirmed cases are sent to the NMHED, but that salmonella and shigella species isolates are sent to the state scientific laboratory which, in turn, reports to the NMHED. The high rates of reporting of laboratory-confirmed cases documented in this study demonstrate that the system is working efficiently.

Connell, *et al*, have described the opportunities and hazards in the use for research of datasets designed and compiled for other purposes.[6] Our study exemplifies such limitations. Although ICD-9-CM code assignments were not sensitive for detection and surveillance of the notifiable infectious diseases we chose for this study, they were congruent with the clinical picture and may have identified potential cases not detected by laboratory-based surveillance. Conditions whose diagnoses rely predominantly on clinical evidence (e.g., injuries) are likely to be more accurately identified by ICD-9-CM code surveillance. Although further studies on the feasibility of inpatient and outpatient data systems for surveillance are needed, access to a dataset combining laboratory, inpatient, and outpatient information holds potential for disease surveillance. □

Acknowledgments

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Physical Activity on the Job and Cancer in Missouri

Ross C. Brownson, PhD, Jian C. Chang, PhD, James R. Davis, MS, and Carol A. Smith, MSPH

ABSTRACT

We conducted a series of case-control studies to investigate the risks of 16 cancer types in relation to occupational physical activity. These studies were based on Missouri Cancer Registry data for 17,147 White male cancer patients registered between 1984 and 1989. Colon cancer risk was increased for both the moderate (odds ratio (OR) = 1.1; 95% confidence interval (CI) = 1.0, 1.3) and low (OR = 1.2; 95% CI = 1.0, 1.5) activity levels. Similar elevations were observed for prostate cancer at the moderate (OR = 1.1; 95% CI = 1.0, 1.3) and low (OR = 1.5; 95% CI = 1.2, 1.8) levels of activity, and for cancer of the testis at the low activity level (OR = 2.2; 95% CI = 1.3, 3.7). An opposite trend ($p < 0.01$) was noted for lung cancer, which showed decreased risk at the moderate (OR = 0.9; 95% CI = 0.8, 1.0) and low (OR = 0.8; 95% CI = 0.6, 0.9) activity levels. These associations suggest that further study of the relationship between physical activity and site-specific cancer risk is warranted. (*Am J Public Health* 1991;81:639–642)

Introduction

A growing body of evidence suggests an inverse relation between physical activity and risk of colon cancer. Recent studies,[1–11] from diverse populations, have identified an association between occupational or recreational physical activity and colon cancer risk, although one report found no such relation.[12] Elevations in risk of colon cancer in relation to low physical activity have typically ranged from 20-100 percent.[1–11]

Few studies have evaluated the association between physical activity and other types of cancer. Recent findings suggest that physical activity may be associated with several cancer types including cancer of the stomach,[8] prostate,[9] and breast.[10–13]

To investigate the risks of various cancer types in relation to occupational physical activity, we conducted a series of case-control studies based on data from a statewide cancer registry.

Methods

Subjects were identified through the Missouri Cancer Registry for the time period January 1984 through May 1989. The Registry is maintained by the Missouri Department of Health and has been collecting data on incident cancer cases from public and private hospitals since 1972. Hospital reporting has been mandated by law since 1984. Reporting procedures and validity issues have been discussed in more detail elsewhere.[14]

The current study involved a series of case-control studies that included White male cancer patients who were 20 years of age or older at the time of diagnosis. Men with cancer of ill-defined and unknown primary sites (*International Classification of Diseases for Oncology*[15] (ICD-O) codes 195 and 199) were excluded. Selection was limited to White males due to the small

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Address reprint requests to Ross C. Brownson, PhD, Director, Division of Chronic Disease Prevention and Health Promotion, Missouri Department of Health, 201 Business Loop 70 West, Columbia, MO 65203. Ms. Smith is also with that Division; Dr. Chang is with the Missouri Cancer Registry; Mr. Davis is with the Bureau of Smoking, Tobacco and Cancer, all with the Missouri Department of Health. This paper, submitted to the Journal June 8, 1990, was revised and accepted for publication October 23, 1990.

TABLE 1—Number of Cancer Patients by Site in Missouri Males, 1984–89

Sites	ICD-O codes: Topography	ICD-O codes: Morphology	Number of Cases
Buccal cavity	140–149		585
Esophagus	150		237
Stomach	151		358
Colon	153		1,838
Rectum	154		812
Pancreas	157		409
Larynx	161		335
Lung	162		4,700
Malignant melanoma		8720/3	143
Prostate	185		2,878
Testis	186		252
Bladder	188		1,080
Kidney/other urinary organs	189		515
Brain/other nervous system	191, 192		332
Non-Hodgkin's lymphoma		*	536
Leukemia		9800/3-9940/3	438
Other cancers			1,699
Total			17,147

*Non-Hodgkin's lymphoma working formulation.[17]

number of other races and the lack of occupational diversity among females.

Data on occupation and tobacco use are routinely abstracted from the hospital medical record at the time of diagnosis for the Missouri Cancer Registry. Hospital registrars record employment information for cancer patients using a standardized protocol. These data are subsequently coded by trained cancer data specialists on the Registry staff using the three-digit 1980 US Census codes.[16] The smoking history of each patient is characterized according to status (i.e., never, former, or current use) and, among current smokers, according to the number of packs smoked per day.

Analysis was limited to 17,147 subjects from the Registry who had codable occupational information. The remaining subjects had no occupational information or noninformative job titles such as "retired."

The classification scheme of Garabrant *et al.*[1] with modifications for Missouri work requirements,[11] was used to categorize occupations according to the level of physical activity. Individual occupations were classified according to whether physical activity is required more than 80 percent of the time (high activity), 20–80 percent of the time (moderate activity), or less than 20 percent of the time (low activity). Subjects were then assigned to one of the three activity levels.

In the analysis for each cancer site, all other cancer registrations formed the control group. For example, in the analysis of the buccal cavity (ICD-O codes 140-149), the case group consisted of 585 patients with this cancer type and the control group consisted of 16,562 "other" cancer patients. The distribution of study subjects by cancer type is shown in Table 1.

Maximum likelihood estimates of the odds ratio (OR)[18] were adjusted for age and smoking, and trend tests were conducted.[19] The highest level of activity was used as the reference category.

In calculating site-specific ORs, colon cancer was excluded from the control group since physical activity has been consistently shown to be a risk factor for colon cancer.[1–11] In addition, because some nonsedentary jobs may be at increased risk of lung cancer due to carcinogenic exposures, ORs for physical activity and lung cancer were adjusted according to *a-priori* job assignments. Occupations were assigned to a high-risk or low-risk category based on a previous study of occupation and lung cancer in Missouri.[20]

Results

Table 2 presents the risks of various cancer types according to the level of occupational physical activity. Elevated colon cancer risk was observed for both the moderate (OR = 1.1; 95% confidence interval (CI) = 1.0, 1.3) and low (OR = 1.2; 95% CI = 1.0, 1.5) levels of activity, with an inverse trend ($p = 0.05$) in risk. Among anatomic subsites of colon cancer, risk was highest for cancer of the cecum, with an OR of 1.3 (95% CI = 0.9, 1.9) at the moderate activity level and an OR of 1.8 (95% CI = 1.0, 3.0) at the low activity level.

Similarly, prostate cancer showed increased risk for moderate (OR = 1.1; 95% CI = 1.0, 1.3) and low (OR = 1.5; 95% = 1.2, 1.8) activity levels. Elevated risk of cancer of the testis was noted for the low activity level (OR = 2.2; 95% CI = 1.3, 3.7). Inverse linear trends were observed for cancers of the prostate ($p < 0.01$) and testis ($p < 0.01$).

An opposite trend ($p < 0.01$) was seen for lung cancer. Decreased risk was shown for moderate (OR = 0.9; 95% CI = 0.8, 1.0) and low (OR = 0.8; 95% CI = 0.6, 0.9) activity levels. Of the major histologic types of lung cancer, the strongest gradient in risk was shown for squamous cell carcinoma. For this cell type, the OR for moderate activity was 1.0 (95% CI = 0.8, 1.1) and the OR for low activity was 0.6 (95% CI = 0.4, 0.8), with a significant gradient in risk ($p < 0.01$).

Discussion

This study corroborates recent studies[1–11] that suggest a relation between low physical activity and colon cancer risk although risk estimates were lower than those reported in some previous studies. Other possible associations were identified between physical activity and cancers of the prostate, testis, and lung.

The physiologic mechanism by which physical activity decreases colon cancer risk is not yet clearly defined but it may relate to increased colon peristalsis[21] and a subsequent decrease in the transit time of the stool and the accompanying exposure of the colon to fecal carcinogens.[4] In addition, experimental animal studies suggest a decreased incidence and growth of induced and transplanted tumors associated with higher levels of physical activity.[22,23]

We found inverse associations between occupational physical activity and cancers of the prostate and testis. Albanes *et al.*[9] recently reported a significant inverse relation between recreational exercise and prostate cancer risk. Previous studies of occupational mortality from Washington State[10] and among San Francisco longshoremen[12] have shown some evidence of a relation between prostate cancer and low job activity. However, other studies[8–12] have shown little evidence of a physical activity-prostate cancer association. The only previous study that investigated physical activity and can-

TABLE 2—Odds Ratios* and 95% Confidence Intervals for Cancer according to Occupational Physical Activity Level in Males, Missouri, 1984–89

Site	Activity Level	Number of Cases	Odds Ratio	95% Confidence Interval	Trend p value
Buccal cavity	moderate	457	1.0	0.8, 1.2	
	low	42	1.1	0.8, 1.7	0.37
Esophagus	moderate	189	1.0	0.7, 1.4	
	low	11	0.7	0.3, 1.4	0.18
Stomach	moderate	281	1.0	0.8, 1.4	
	low	28	1.4	0.9, 2.2	0.13
Colon	moderate	1,476	1.1	1.0, 1.3	
	low	120	1.2	1.0, 1.5	0.05
Rectum	moderate	646	1.1	0.9, 1.3	
	low	55	1.2	0.8, 1.7	0.16
Pancreas	moderate	339	1.3	1.0, 1.8	
	low	22	1.1	0.6, 1.9	0.17
Larynx	moderate	274	1.0	0.7, 1.4	
	low	11	0.5	0.3, 1.0	0.08
Lung	moderate	3,687	0.9	0.8. 1.0	
	low	230	0.8	0.6, 0.9	<0.01
Malignant melanoma	moderate	111	1.0	0.6, 1.6	
	low	11	1.2	0.5, 2.6	0.39
Prostate	moderate	2,301	1.1	1.0, 1.3	
	low	200	1.5	1.2, 1.8	<0.01
Testis	moderate	189	1.1	0.8, 1.7	
	low	31	2.2	1.3, 3.7	<0.01
Bladder	moderate	565	1.1	0.9, 1.3	
	low	66	1.1	0.8, 1.5	0.28
Kidney/other urinary organs	moderate	413	1.2	0.9, 1.5	
	low	36	1.3	0.9, 2.0	0.10
Brain/other nervous system	moderate	256	1.0	0.7, 1.3	
	low	27	1.3	0.8, 2.1	0.27
Non-Hodgkin's lymphoma	moderate	425	1.0	0.8, 1.4	
	low	37	1.2	0.8, 1.8	0.23
Leukemia	moderate	346	1.0	0.8, 1.3	
	low	28	1.1	0.7, 1.7	0.47
Other cancers	moderate	1,384	1.2	1.0, 1.4	
	low	95	1.0	0.8, 1.3	0.24

*Adjusted for age, smoking, and carcinogenic job assignment (lung cancer only); the high activity level was the reference category.

cer of the testis was the study of Harvard and University of Pennsylvania alumni[12] in which no association was observed.

We found a positive association between occupational physical activity and lung cancer risk, although uncontrolled confounding for smoking habits may be partially responsible for this relationship. Earlier studies[8,9,12] have shown a general tendency toward decreased lung cancer risks associated with various measures of higher physical activity levels, although findings have not been entirely consistent.

Severson *et al.*[8] recently reported an increased risk of stomach cancer in relation to several measures of physical activity. These Missouri data showed no evidence of a physical activity-stomach cancer association. We found a slight association between low physical activity and rectal cancer risk, which has been reported inconsistently in previous reports.[1,2,8–10,12]

Our study has several limitations. We used occupational information collected in conjunction with cancer incidence reporting to produce a surrogate measure of physical activity. However, a recent study[14] that compared risk factor information contained in the Registry record with that obtained by interview found reasonable agreement for occupational data. We were also unable to adjust for some potential confounders such as socioeconomic status and dietary history. Although we lacked information on avocational physical activity, it has been reported that the majority of total physical activity in American workers occurs on the job.[24]

Major strengths of our study are the greater accuracy in hospital-reported tumor diagnosis, compared with death certificates,[25] and the ability to analyze data by histologic type or by anatomic subsite.

Our findings for cancers of the prostate, testis, and lung should be considered preliminary and require confirmation in other studies with more detailed measurements of physical activity and potential confounders. ☐

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Health Care Received by American Children: NCHS Report

The vast majority (83 percent) of American children are covered by health insurance, have a regular source of health care (88 percent, and visited a doctor in the past year for routine health care (64 percent) according to a recent report issued by the National Center for Health Statistics. The report cites new evidence, however, that America's minority and poor children receive less care, including preventive services.

The report—*Health Insurance and Medical Care: Health of Our Nation's Children, United States 1988*—is derived from data from the 1988 National Health Interview Survey on Child Health, conducted by NCHS. Questions on health insurance and sources of medical care were requested for all 17,110 children in a nationally representative sample of children 17 years of age and under.

The survey shows that Black newborns and infants were two to three times more likely to have lacked a routine visit for a check-up or any immunization during the first year of life than were White newborns and infants. Hispanic children were twice as likely to have no health insurance coverage, either public or private, as non-Hispanics.

Most children with a regular source of health care had private care; however, 16 percent of those children (8.7 million) received health care from hospital clinics, emergency rooms, walk-in or emergency centers or other clinics. Clinics were the source of care for 47 percent of Black infants, compared to 16 percent of White infants. For Hispanic infants, almost half received health care in clinics.

The report, *Health Insurance and Medical Care: Health of Our Nation's Children, United States, 1988*—is available from the National Center for Health Statistics, Centers for Disease Control, US Public Health Service, 6525 Belcrest Road, Hyattsville, MD 20782; (301) 436-8500.

Vietnam Era and Vietnam Combat Veterans among the Homeless

Robert Rosenheck, MD, Peggy Gallup, PhD, and Catherine A. Leda, MPH, MSN

ABSTRACT

Of 10,524 homeless veterans assessed in a 43-site VA program, 50 percent served during the Vietnam War era, compared to only 29 percent of all veterans in the general population. This reflects the greater risk of homelessness among men aged 30-44 rather than the impact of Vietnam Era service. The proportion of homeless veterans who served in the Vietnam Theater (44.9 percent), and the proportion exposed to combat fire (40.5 percent) were similar to those of nonhomeless veterans. Homeless combat veterans who are not White were more likely to have psychiatric, alcohol, and medical problems than homeless noncombat Vietnam veterans who are not White. (*Am J Public Health* 1991;81:643–646)

Introduction

Surveys of the homeless have indicated that 43 percent to 50 percent of homeless male veterans served in the military during the Vietnam Era (August, 5, 1964-May 7, 1975),[1-3] compared to only 29 percent of veterans in the general population.[4] During the 1980s, reports of high percentages of Vietnam Era veterans among the homeless appeared virtually simultaneously with a growing body of literature reporting significant post-war readjustment problems among Vietnam combat veterans[5,6] and suggested that veterans who served in Vietnam may be at particularly high risk for homelessness.

Three important shortcomings of these initial studies were: 1) the numbers of Vietnam Era veterans studied were small (53-88 per study); 2) the potentially confounding influence of age was not addressed (Vietnam Era veterans fall into the age cohort that is most at risk for homelessness)[7] and 3) data were not presented on exposure to the presumed combat stressors, whether service in the Vietnam theater of war or exposure to life-threatening combat fire.

In 1987-88, the Department of Veterans Affairs (VA) Homeless Chronically Mentally Ill Veterans Program conducted intake assessments on 10,524 homeless veterans, of whom 5,287 (50 percent) reported service during the Vietnam Era. In this paper, demographic, military service and clinical assessment data on these veterans, and published data on Vietnam Era veterans in the general population, are used to examine: 1) whether Vietnam Era veterans are disproportionately represented among homeless veterans, when age is taken into consideration; 2) whether homeless Vietnam Era veterans are more likely to have served in Vietnam or to have been exposed to combat fire in Vietnam than Vietnam Era veterans in the general population; and 3) whether homeless veterans who served in Vietnam or who were exposed to combat fire there have more frequent medical, psychiatric or substance abuse problems as compared to a reference group of homeless veterans who did not serve in Vietnam.

Methods

Subjects

The Homeless Chronically Mentally Ill (HCMI) Veterans Program was implemented in May 1987, at 43 VA Medical Centers across the country. Sites were selected on the basis of the estimated number of homeless persons in their service area, their capacity to conduct outreach to the homeless and the presence of a coalition of agencies assisting the homeless in their area.

Homeless veterans were contacted through several routes: 52.7 percent were contacted through community outreach efforts in which social workers and nurses visited shelters, soup kitchens and other places frequented by the homeless, seeking out veterans in need of assistance; 26.1 percent came to VA facilities on their own; 13.3 percent were referred by non-VA programs for the homeless; and 7.9 percent came through unspecified routes. Veterans were offered intake assessments if they were judged to be eligible for VA benefits, lacked a conventional place to sleep, and expressed interest in VA health care services or other benefits.

Procedures

A standardized intake form recording data gathered through a structured interview, and clinicians' observations of major psychiatric symptoms, was completed on veterans assessed during the first 10.5 months of program operation (May 15, 1987-March 30, 1988). The interview addressed self-reported age, gender, ethnicity, duration of the current episode of homelessness, dates of admission and discharge from military service, receipt of combat pay, exposure to fire during com-

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From the Northeast Program Evaluation Center, VA Medical Center, West Haven, CT (all authors); and the Yale University Department of Psychiatry (Rosenheck). Address reprint requests to Robert Rosenheck, MD/182, Northeast Program Evaluation Center, Veterans Administration Medical Center, 950 Campbell Avenue, West Haven, CT 06516. This paper, submitted to the Journal May 2, 1990, was revised and accepted for publication September 19, 1990.

TABLE 1—Percent of Veterans in Five-year Age Cohorts: HCMI Veterans Program (May 15, 1987–March 31, 1988) and All US Veterans, 1987

Age Cohort (Years)	HCMI Veterans Program N = 10,193 +	All U.S. Veterans* N = 28,688,000
	%	%
<25	1.4	1.1
25–29	7.9	3.9
30–34	16.9	5.7
35–39	20.9	8.3
40–44	16.5	12.0
45–49	9.1	8.7
50–54	8.6	9.4
55–59	8.4	12.4
60–64	5.8	14.4
65–69	2.7	12.9
70–74	1.1	6.9
>75	0.5	4.3
Total	100.0	100.0

*VA Annual Report, 1989.
+data missing on 331 veterans.

bat service, and past history of psychiatric and substance abuse hospitalization. Alcohol and drug abuse problems were identified through a combination of patient self-reports (e.g., "Do you have a problem with alcohol dependency now?") and clinician observations. Major psychiatric problems were assessed with a 10-item summed Psychiatric Problem Index (PPI), consisting of three self-report items and seven clinical observations, (Cronbach's alpha = 0.71). On a sub-sample of 1,318 veterans asssessed in this Program, a PPI of two or more was associated with a major DSM-III psychiatric disorder (schizophrenia or other psychosis, bipolar disorder or depressive disorder), as determined by a psychiatrist's examination, with a sensitivity of 80.1 percent and a specificity of 58.0 percent. A score of two or more on the PPI was thus considered evidence of significant psychiatric disorder.

Military service during the Vietnam Era was categorized as follows. All veterans who served between August 7, 1964 and May 5, 1975 are officially classified by VA as *Vietnam Era* veterans. Within this group, *Vietnam Era Only* veterans (those who did not serve in Vietnam, and who were not exposed to warzone stress) were distinguished from *Vietnam Theater* veterans. Within the group of Vietnam Theater veterans, *noncombat exposed* veterans (those who served in Vietnam but were not under fire) were distinguished from *combat exposed* veterans.

Published data on veterans in the general population were used for comparison purposes and included the Census Bureau's 1987 civilian population estimates[4] and the Third Survey of Veterans,[8] a 1987 survey of a national sample of 9,442 veterans.

Comparisons were made between the proportions of homeless veterans in various age and military service categories and estimated proportions of veterans in the general population in those categories. In addition, health status indicators for noncombat-exposed and combat-exposed Vietnam Theater veterans were compared with health status indicators for veterans who did not serve in Vietnam. Measurement precision was expressed as the 95 percent confidence interval of differences in percentages.

Results

Veteran Characteristics

Age Distribution: All Service Eras: The median age among *all* veterans in the HCMI Program was 40.1 years; 54.3 percent were 30-44 years old (Table 1). Homeless veterans in this program were thus similar in age to other homeless Americans who have a median age of 36.1 years.[7] In contrast, the median age among veterans in the general veteran population in 1987 was 54.4 years; only 26 percent were 30-44 years old. Among veterans, as in the general population,[7] the largest percentages of the homeless are in their 30s and early 40s.

Vietnam Era Service: Over half of all veterans in the HCMI Veterans Program served during the Vietnam Era, compared to only 29 percent in the general veteran population. However, within matched five-year age cohorts, the percentages of Vietnam Era veterans in the HCMI Veterans Program were generally similar to those in the general veteran population, although among the homeless, greater percentages of Vietnam Era veterans were found within younger age cohorts (Table 2). Among both homeless veterans and veterans in the general population, over 80 percent of 30-44 year-olds were Vietnam Era veterans (87 percent in the HCMI Veterans Program; 81 percent in the general veteran population).

Characteristics of Vietnam Era Veterans: Vietnam Era veterans were 99 percent male; 54.6 percent White; 36.8 percent Black; 6.9 percent Hispanic; and 1.7 percent other. The racial distribution was similar to that reported in other studies of the homeless.[7] The duration of the current episode of homelessness was less than one month for 25.8 percent, from one month to one year for 43.9 percent and more than one year for 30.3 percent.

Vietnam Theater Service: Among Vietnam Era veterans assessed by the HCMI Veterans Program 44.9 percent reported service in the Vietnam Theater; 40.5 percent reported exposure to combat and 4.4 percent reported no combat exposure. These proportions are quite similar to those of Vietnam Era veterans in the general population, among whom 48.7 percent report service in the Vietnam Theater; 38.4 percent report exposure to combat and 10.3 percent report war zone service but no combat exposure.[8]

The percentage of White homeless Vietnam Era veterans who reported service in the Vietnam Theater was somewhat greater than that of other racial-ethnic groups (46.2 percent vs 42.4 percent). In contrast, in the general population of veterans 48.0 percent of Whites report service in the Vietnam Theater compared to 52.1 percent of other racial-ethnic groups.

Clinical Assessment of Homeless Vietnam Era Veterans: Among homeless Vietnam Theater combat veterans who were not white, current medical, psychiatric and alcohol problems, as well as past histories of psychiatric hospitalization, were more common than among noncombat homeless veterans of similar race-ethnicity (Table 3). While the same trends were apparent among Whites, they were of lesser magnitude.

TABLE 2—Proportion of Veterans Who Served During the Vietnam Era by Five-year Age Cohorts: HCMI Veterans Program (May 15, 1987–March 31, 1988) and All US Veterans, 1987

Age Cohort (Years)	HCMI Veterans Program N = 10,193	All US Veterans N = 28,688,000*
	%	%
<25	2.1	0.0
25–29	2.5	1.1
30–34	66.4	42.0
35–39	95.6	89.9
40–44	91.4	94.3
45–49	40.8	49.6
50–54	6.1	13.7
55–59	2.6	7.9
60–64	0.8	3.6
65–69	0.4	2.4
70–74	0.9	1.8
>75	1.9	1.1
Total (All ages)	51.0	28.9

*VA Annual Report, 1989.
+Data missing on 331 veterans.

TABLE 3—Comparison of Clinical Problems Identified on Screening Assessment between Veterans Who Did Not Serve in Vietnam and Those Exposed to Combat, by Race

	Vietnam Era Only	Vietnam Theater/ Combat Exposed	Difference in Percent (95% CI)
White	N = 1,528	N = 1,189	
	%	%	%
Medical Problems	41.0	44.6	3.5 (−0.2,7.3)
Psychiatric Problem	50.0	52.1	2.1 (−1.7,5.9)
Alcohol Problems	51.2	52.8	1.6 (−2.2,5.4)
Drug Problems	17.1	18.4	1.3 (−1.6,4.2)
Past Psychiatric Hospital	41.5	42.5	1.1 (−2.7,4.8)
Past Alcohol Hospital	50.2	48.9	−1.3 (−5.1,2.5)
Past Drug Hospital	17.7	20.2	2.5 (−0.4,5.5)
Others	N = 1,387	N = 952	
	%	%	%
Medical Problems	37.9	45.8	7.9 (3.7,12.0)
Psychiatric Problem	38.0	48.4	10.5 (6.3,14.6)
Alcohol Problems	44.9	49.1	4.2 (0.0,8.4)
Drug Problems	26.7	26.9	0.2 (−3.5,3.9)
Past Psychiatric Hospital	28.9	35.5	6.6 (2.6,10.5)
Past Alcohol Hospital	38.6	42.4	3.7 (−0.4,7.8)
Past Drug Hospital	26.0	26.3	0.3 (−3.4,4.0)

Discussion

Several methodological limitations of this study require comment. Although the data were collected in a national program, they were not gathered through a systemic sampling procedure and findings cannot be confidently generalized to the national population of homeless Vietnam Era veterans. Both sampling and reporting biases may influence our findings. First, data were collected through a VA clinical program, in which an over-representation of combat-stressed veterans might be expected. Second, because the data were gathered in a VA program, veterans may have tended to overreport their combat experiences, assuming that such reports would improve their access to VA health care services and financial benefits. On the other hand, many Vietnam veterans have been critical of VA services,[9] and some have been reluctant to use them,[10] possibly biasing the representation of Vietnam Theater veterans downward. In our view, Vietnam service is more likely to have been overcounted in this study than undercounted.

The data presented suggest that the large percentage of Vietnam Era veterans among homeless veterans is best explained by the specific vulnerability of men in their 30s and early 40s to homelessness.[7] Neither Vietnam Era service, Vietnam Theater service, nor exposure to combat fire appear to be unexpectedly common among homeless veterans studied.

This conclusion is consistent with findings from the National Vietnam Veterans Readjustment Study (NVVRS),[6] conducted in 1987. The NVVRS did note, however, that veterans who met diagnostic criteria for Posttraumatic Stress Disorder (15.2 percent of all Vietnam Theater veterans) were far more likely to have been homeless than those who did not. While Vietnam service, by itself, does not appear to be a major risk factor for homelessness, those veterans who were most affected by combat may be at greater risk for homelessness than others.

Vietnam veterans in the HCMI Veterans Program who were not White and who were exposed to combat have more frequent psychiatric, medical and alcohol problems than those veterans who are not White who did not serve in Vietnam. These findings are consistent with those of the NVVRS which found substantial differences in health status measures between veterans who were exposed to high levels of combat and those not exposed, and a considerably higher prevalence of Posttraumatic Stress Disorder among those who are not White than among Whites.[6] The absence of significant differences in health status measures among Whites may reflect the limited specificity of our screening instruments, although further research is needed to clarify the meaning of these findings. While not overrepresented among homeless veterans, homeless Vietnam combat veterans, like Vietnam combat veterans examined in other studies, appear to face more frequent medical and psychiatric problems and may require special services. ☐

Acknowledgments

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Determinants of Late Stage Diagnosis of Breast and Cervical Cancer: The Impact of Age, Race, Social Class, and Hospital Type

Jeanne Mandelblatt, MD, Howard Andrews, PhD, Jon Kerner, PhD, Ann Zauber, PhD, and William Burnett, MD

ABSTRACT

Previous studies of the relationship between cancer stage, age, and race have not controlled for social class and health care setting. Logistic regression analyses, using information from the New York State Tumor Registry and area-level social class indicators, demonstrated that, in New York City, older Black, lower class women in public hospitals were 3.75 and 2.54 times more likely to have late stage breast or cervical cancer, respectively, than were younger White, high social class women in non-public hospitals. (*Am J Public Health* 1991;81:646–649)

Introduction

Age and race have each individually been associated with cancer stage.[1–5] However, previous studies have not controlled for the effects of social class and health care setting. This paper presents findings from logistic regression analyses to quantify the individual and combined effects of age, race, socioeconomic class, and type of health care setting on breast and cervical cancer stage.

Methods

Cancer Data-Tumor Registry

The study sample consisted of all cases of breast and cervical cancer among New York City (NYC) residents that were reported to the New York State Department of Health Tumor Registry between 1980 and 1985. More than 90 percent of incident cases are reported to the registry from NYC.[6] In the study period, there were 22,111 breast cancer, 2,930 invasive cervical cancer, and 6,408 cervical carcinoma in-situ cases with known stages reported. Stage was unknown for 7.5 percent and 2.6 percent of the breast and cervical cases, respectively. Missing stage was associated with age ($p < .0001$) and missing marital status ($p < .01$) for both cancers. Age was classified into five-year categories. Race was reported as either Black or White. There were insufficient data to include Hispanic origin in the analysis. Marital status was dichotomized as "ever" versus "never" married. Hospital type was classed as "public" for the 11 NYC municipal hospitals and "non-public" for the remaining facilities.

Demographic Variables—Census Data

Because individual information on socioeconomic class is not available in registry data, 1980 United States census

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Address reprint requests to Jeanne Mandelblatt, MD, Assistant Professor, Division of Cancer Control, Department of Epidemiology and Biostatistics, Memorial Sloan-Kettering Cancer Center, Box 60, 1275 York Avenue, New York, NY 10021. Drs. Kerner and Zauber are also with that Department at the Cancer Center; Dr. Andrews is with New York State Psychiatric Institute, New York City; Dr. Burnett is Director, Tumor Registry, New York State Department of Health, Albany. This paper, submitted to the Journal July 12, 1990, was revised and accepted for publication December 5, 1990.

TABLE 1—Sample Characteristics for Cases Used in Each Model

Characteristics	Model 1 Breast Late vs Early	Model 2 Breast Distant vs Regional	Model 3 Cervical Late vs Early	Model 4 Cervical Invasive vs In-Situ
Number of Cases[a]	19,132	9,991	2,343	7,182
State (% Late)	52.2	NA	45.1	32.6
Age in years, mean (S.D.)	79.5	60.6 (13.92)	52.4 (15.40)	41.7 (15.8)
Race (% Black)	17.6	20.0	41.0	36.8
Hospital (% Public)	6.1	7.7	16.1	14.0
% Ever Married	84.5	84.3	80.4	70.5
% High School[c], mean (S.D.)	62.3 (16.4)	61.0 (16.2)	54.8 (16.2)	57.3 (17.2)
Median Family Income[c], mean (S.D.)	$16,139 (6,427)	$15,756 (6,333)	$12,966 (5,746)	$13,796 (6,190)

a) The number of cases having complete data for all variables.
b) Percent post-menopausal.
c) These variables were added to the NYS Tumor Registry data set using values for the census tract in which the case lived.

TABLE 2—Summary of Logistic Regression Analysis: Prediction of Late Stage Breast Cancer

	Late vs Early		Distant vs Regional	
Variables	Regression Coefficient	Odds Ratio (95% Confidence Interval) (p-value)	Regression Coefficient	Odds Ratio (95% Confidence Interval) (p-value)
Age[a]	.086	1.09 (1.02, 1.17) (p < .015)	.086	1.09 (1.08, 1.11) (p < .001)
Race (1 = White, 2 = Black)	.215	1.24 (1.14, 1.35) (p < .001)	−.083	.92 —[c] (p < .15)
Hospital Setting (0 = non-public, 1 = public)	.412	1.51 (1.32, 1.71) (p < .001)	.322	1.38 (1.18, 1.62) (p < .001)
Marital Status (0 = never, 1 = ever)	−.041	.96 —[c] (p < .369)	−.357	.70 (.63, .79) (p < .001)
Income[b] ($5,000 increments)	.029	1.03 —[c] (p < .055)	−.020	.98 —[c] (p < .482)
Education[b] (% high school, 5% increments)	−.051	.95 (.94, .96) (p < .001)	−.030	.97 (.95, .99) (p < .002)
Intercept	.29		−.96	

a) All age in 5-year intervals, except late vs early breast, which is post vs pre-menopausal.
b) Note: These are continuous variables and the odds ratios represent the increased risk for each incremental increase in the variable, compared to the next lowest level of the variable.
c) Confidence interval includes one.

data on the median family income and the percent of adults with a high school education in the census tract of residence of the case were used as social class measures.

Logistic Regression Models

Two models were developed for each cancer based on groupings with similar treatment patterns. For breast cancer, the first model predicted the outcome of late stage (regional and distant) compared to early stage (localized and in-situ). As a result of recent advances in the treatment of regional disease, a second model was developed comparing distant to regional disease. For cervical cancer, the outcome of the first model was late stage (regional and distant) versus early stage (localized) invasive cancer. In addition, invasive and in-situ disease were compared in a second model as a measure of the success of Pap screening. For all models the independent variables were age, race, marital status, hospital type, income, and education. Logistic regression models were developed using SAS programs. Table 1 presents summary statistics for the cases included in the models.

Results

The logistic regression analysis indicated that, controlling for the other variables, the following factors were associated with late stage breast cancer: postmenopausal age, Black race, low education, and public hospital use (Table 2). Lower income was also weakly associated with late stage ($p < .055$). In the second model, age, public hospital use, marital status, and low education were each associated with an increased risk of distant disease.

Regression coefficients and intercepts from each model were used to calculate the odds of late stage disease for combinations of variables (Table 3). For example, the odds of having distant, compared to regional, breast cancer were 4.77 times higher for a Black woman who was age 70 to 74, unmarried, in the lowest income and education groups, and who used the public hospital system, compared to a White married woman age 40 to 44, in the highest income and education groups, using a non-public hospital.

For invasive cervical cancer, age and public hospital setting were each associated with an increased risk of late stage disease (Table 4). In the second model, age, race, hospital setting, income, and education were each associated with invasive disease.

The composite odds of a 70 to 74 year old, unmarried Black woman in the lowest income and educational group who used the public hospital system having late stage invasive cervical cancer were 2.54 times greater than for a White, married, 40 to 44 year old woman in the highest income and educational levels and using a non-public hospital (Table 5). The odds for

TABLE 3—Risk of Late Stage Breast Cancer for Women of Differing Demographic Profiles

	Late vs Early[a]	Distant vs Regional[b]
	Odds Ratio[c]	Odds Ratio[c]
Profile 1: Base Case[d]		
70–74 years, Black, poor, low education, public hospital, unmarried	3.75	4.77
Profile 2: Young		
40–44 years, Black, poor, low education, public hospital, unmarried	3.43	2.82
Profile 3: White		
70–74 years, White, poor, low education, public hospital, unmarried	3.02	not significant[e]
Profile 4: Non-public hospital		
70–74 years, Black, poor, low education, non-public hospital, unmarried	2.49	3.46
Profile 5: 90% High School		
70–74 years, Black, poor, high education, non-public hospital, unmarried	2.04	3.33

a) The minimum odds for the lowest risk case (a pre-menopausal, White, married woman using a non-public hospital and living in a community where 90% of the adult population completed high school and the median income is $40,000 per year) is .67.

b) The minimum odds for the lowest risk woman (a 40–44 year old, White, married woman using a non-public hospital and living in a community where 90% of the adult population completed high school and the median income is $40,000 per year) is .35.

c) Odds ratios are computed from the division of the maximum odds by the minimum odds. Minimum and maximum odds are calculated using regression coefficients from the logistic models.

d) Low education and education represent living in a community where 30% of adults completed high school and the median income is $10,000 per year.

e) The variable of interest was not significant in the original regression model.

a woman of the same profile having invasive compared to in-situ disease were 39.75 times higher than for her younger, White, non-disadvantaged counterpart.

Discussion

This study demonstrates that elderly, Black, lower socioeconomic status (SES) women who use public hospitals are at extremely high risk of having their breast and cervical cancers diagnosed at late stages. While our findings are consistent with and extend the findings of prior researchers,[1-5] there are two caveats that should be considered in evaluating our conclusions.

First, in our analysis missing stage was associated with age. However, since older women with advanced disease are less likely to be staged than younger women,[7,8] the missing stage information is likely to underestimate the observed effect of age on stage.

Second, since the NY State Tumor Registry does not contain social class indicators, we assigned to each woman the median income and educational level of her community of residence. While we have no way to determine the relationship of the individual and community SES indicators, several studies of cancer incidence and/or mortality have used group-level data to define socioeconomic status[9-15] and have drawn similar conclusions to studies that use individual information.[16-19] Group data can also capture characteristics of the individual's social "context" which may provide information about class gradients that may not be reflected by individual characteristics.[20,21] It is possible that hospital type is serving as such a "contextual" measure in our analysis.

Our findings suggest that cancer control interventions targeted to elderly disadvantaged women, particularly those of color and those who use public hospitals, have the potential to decrease avoidable cancer morbidity and mortality. ☐

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TABLE 4—Summary of Logistic Regression Analysis: Prediction of Late Stage Cervical Cancer

	Late vs Early		Invasive vs In-Situ	
Variables	Regression Coefficient	Odds Ratio (95% Confidence Interval) (p-value)	Regression Coefficient	Odds Ratio (95% Confidence Interval) (p-value)
Age[a] (5 year increments)	.113	1.12 (1.09, 1.15) (p < .001)	.358	1.43 (1.40, 1.46) (p < .001)
Race (1 = White)	.113	1.12 —[b] (p < .118)	.148	1.16 (1.02, 1.31) (p < .020)
Hospital Setting (0 = non-public, 1 = public)	.247	1.28 (1.02, 1. 0) (p < .035)	.223	1.25 (1.07, 1.48) (p < .006)
Marital Status (0 = never, 1 = ever)	−.128	.88 —[b] (p < .256)	.104	1.11 —[b] (p < .144)
Income[a] ($5,000 increments)	.000	1.00 —[b] (p < .927)	−.128	.88 (.81, .95) (p < .022)
Education[a] (% high school, 5% increments)	.000	1.00 —[b] (p < .945)	−.030	.97 (.94, .97) (p < .001)
Intercept		−1.66		−.382

a) Note: These are continuous variables and the odds ratios represent the increased risk for each level compared to the next lowest level.

b) Confidence interval includes one.

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TABLE 5—Risk of Late Stage Cervical Cancer for Women of Differing Demographic Profiles

	Late vs Early[a]	Invasive vs In-Situ[b]
	Odds Ratio[e]	Odds Ratio[e]
Profile 1: Base Case[c] 70–74 years, Black, poor, low education, public hospital, unmarried	2.54	39.75
Profile 2: Young 40–44 years, Black, poor, low education, public hospital, unmarried	1.28	4.6
Profile 3: White 70–74 years, White, poor, low education, public hospital, unmarried	not significant[d]	34.33
Profile 4: Non-public hospital 70–74 years, Black, poor, low education, non-public hospital, unmarried	1.99	31.69
Profile 5: 90% High School 70–74 years, Black, poor, high education, non-public hospital, unmarried	not significant	27.40
Profile 6: $40,000 income 70–74 years, Black, high income, low education, non-public hospital, unmarried	not significant	18.16

a) The minimum odds for the lowest risk case (a pre-menopausal, White, married woman using a non-public hospital and living in a community where 90% of the adult population completed high school and the median income is $40,000 per year) is .67.

b) The minimum odds for the lowest risk woman (a 40–44 year old, White, married woman using a non-public hospital and living in a community where 90% of the adult population completed high school and the median income is $40,000 per year) is .35.

c) Low education and education represent living in a community where 30% of adults completed high school and the median income is $10,000 per year.

d) The variable of interest was not significant in the original regression model.

e) Odds ratios are computed from the division of the maximum odds by the minimum odds. Minimum and maximum odds are calculated using regression coefficients from the logistic models.

Public Health and the Law

The Effect of Mandatory Premarital HIV Testing on Marriage: The Case of Illinois

Background

In the past 10 years, AIDS (acquired immunodeficiency syndrome) has become a dominant public health concern. During the late 1980s, state governments took steps to counter the health threat of AIDS. One of the most controversial was the requirement of the State of Illinois that applicants for marriage licenses be tested for human immunodeficiency virus (HIV) antibodies. Among the states, Illinois and Louisiana adopted mandatory premarital testing; only in Illinois was a law implemented for a significant length of time.[1]

Effective January 1, 1988, the Illinois marriage law was amended to require a physician's statement on HIV testing be submitted to the county clerk before a marriage license could be issued. The statement affirmed that both parties to the marriage were tested for HIV antibodies and advised of the test results. The HIV testing requirement was one of a number of AIDS-related measures adopted by the state legislature. In this case, the feeling that people getting married ought to know about HIV infection and the absence of a direct cost to the state was persuasive to lawmakers.[2] The HIV testing requirement was repealed in September 1989.

At the time it was introduced, premarital testing was controversial. Mandatory testing was opposed by the Centers for Disease Control (CDC) and various medical groups.[3,4] Simulation indicated that mandatory premarital testing would be very expensive[5] and probably no more cost-effective for case identification than testing for syphilis.[6]

Upon implementation, the HIV testing requirement was said to depress the frequency of marriages in Illinois. Media reports indicated that couples were crossing state lines to avoid the test,[7] although some observers thought the drop in marriages might be permanent:

> It is ironic that the Illinois screening program has discouraged many people from entering into a legal commitment that fosters a mutual monogamous relationship.[8,9]

This report investigates the hypotheses: (a) that mandatory premarital HIV testing decreased marriages in Illinois; and (b) that the requirement increased crossover marriages in states bordering Illinois. Additional analyses investigated the claim that the frequency of marriages in Illinois has rebounded now that premarital testing is no longer mandatory.[10]

Methods

Illinois Premarital Testing Law

For roughly 20 months, from January 1, 1988 through September 10, 1989, Illinois marriage law required a physician's statement that both parties to the proposed marriage were tested for HIV antibodies and advised of the test results be submitted to the county clerk before a marriage license could be issued.[11] The tests must have been administered within 30 days of the license application. The license was good for 60 days. Positive results on the initial enzyme-linked immunosorbent assay (ELISA) test were to be followed by the Western Blot Assay. A person was considered seropositive for HIV virus if he or she was positive on both tests. Seropositive results were to be reported to the Illinois Department of Public Health.[12]

Table 1 presents marriage license requirements for the seven states included in this study at the time the Illinois premarital testing law became effective.[13] Illinois, like two of the bordering and two of the comparison states, required a blood test for syphilis. In Illinois the marriage license became effective the next day; all the other states had longer waiting periods, although in two bordering and two comparison states, these periods could be waived.

TABLE 1—Marriage License Requirements in Seven States, 1987

State	Blood Test	Waiting Period
Illinois	Yes	Marry next day
Indiana	Yes	72 hours
Wisconsin	Yes	5 days, waivable
Missouri	No	3 days, waivable
Ohio	Yes	5 days
Minnesota	No	5 days, waivable
Kansas	Yes	3 days, waivable

Marriage Time Series

Monthly marriage times series were investigated for seven states for the period from January 1980 through August 1989[14]: (a) Illinois; (b) Indiana, Wisconsin, and Missouri, the three states bordering Illinois largest MSAs (metropolitan statistical areas) that would be expected to show crossover increases; and (c) Ohio, Minnesota, and Kansas, three comparison states

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Address reprint requests to Jack McKillip, PhD, Professor of Psychology, Department of Psychology, Southern Illinois University at Carbondale, Carbondale, IL 62901. This paper was accepted for publication on August 30, 1990 by George J. Annas, JD, MPH, editor of the Public Health & the Law column, and Professor of Health Law, Boston University, School of Public Health, 80 East Concord Street, Boston, MA 02118.

that do not border Illinois and were not expected to show crossover effects. The beginning of the study period was chosen to allow a long enough pre-HIV-testing series to estimate cyclical trends. August 1989 was selected as the final observation because the law was repealed in mid-September 1989.

Monthly marriage frequencies are collected by state health officials and reported to the National Center for Health Statistics (NCHS) which publishes them on a monthly, "provisional" basis. These time series are labeled "provisional" because non-regular reporting schedules across the states require occasional adjustments to published data. During the first half of the study period, these adjustments were noted in the reports. This practice has not been maintained. Marriages are reported by state of occurrence, not by the state in which couples reside.

Primary NCHS interest is in aggregate data, not in maintaining monthly series on a state-by-state basis. Experience indicates that adjustments more likely reflect a random pattern than consistent practice in any of the states.[15] This type of variation in the time series represents a threat of Type II (unreliability) rather than Type I error (history or instrumentation) for the present study. Because of the statistical analysis procedures, monthly time series were used rather than the more stable but less diagnostic yearly time series.

Statistical Analyses

Box-Jenkins ARIMA interrupted times-series analysis[16] was used because of its ability to estimate cyclical trends and to provide unbiased statistical tests of the impact of the premarital testing law. Using this technique, trend and serial dependency parameters first were estimated for the 96 pre-HIV-testing months, January 1980 to December 1987, for each of the seven states. Next an interruption model is tested for the full time series (including the 20 months from January 1988 to August 1989 when HIV testing was required in Illinois). In addition to trend and serial dependency components, the interruption model includes a binary time series (e.g., 000011111) with 0 values for months during the pre-HIV-testing period and with 1 values for testing months. This Testing Law parameter estimates the size and form (if any) of change in the marriage time series at the point when mandatory premarital HIV testing was introduced in Illinois.

Both permanent/temporary, and abrupt/gradual interruption models were tested for all seven states. Only the permanent, abrupt model results are reported (where appropriate) because only this model yielded statistically significant and appropriate testing law parameters. A permanent, abrupt interruption model shifts the level of the time series only at the point of intervention. This shift is maintained for the duration of the intervention.

Three measures of goodness of fit are available for testing models in Box-Jenkins ARIMA times-series analysis,[16] two of which have counterparts in normal multiple regression analysis. Individual parameters are tested for statistical significance by t-tests and overall R^2 is computed for the full models. This R^2 value is generally much larger for times series analysis than is typical for normal multiple regression analysis. A chi-square test is also made of the residual autocorrelation function. This test estimates whether reliable variation remains among the time series observations. Statistical analyses used the 1984 version of the BMDP 2T Box-Jenkins Times Series program.[17]

Results

Decrease of Marriages in Illinois

Figure 1 presents the raw marriage time series for Illinois. Note the recurrent seasonal cycle, marriages up in Summer and down in Winter. The first hypothesis asked whether the time series moved down as a result of premarital HIV testing. Inspecting Figure 1, this hypothesis expects that marriage frequency is lower starting January 1988 than before this date.

The top row of Table 2 presents the results of statistical analyses relevant to the hypothesized drop in marriages in Illinois. The statistically significant seasonal trend component indicates that, over the 10 years studied, marriages in Illinois decreased by 241 a month on a yearly basis (e.g., January 1980 to January 1981). The model also has statistically significant moving average and autoregressive components that are not presented. The statistically significant testing law parameter reveals an abrupt decrease of 1,043 marriages in Illinois in January 1988 when premarital HIV testing became mandatory.

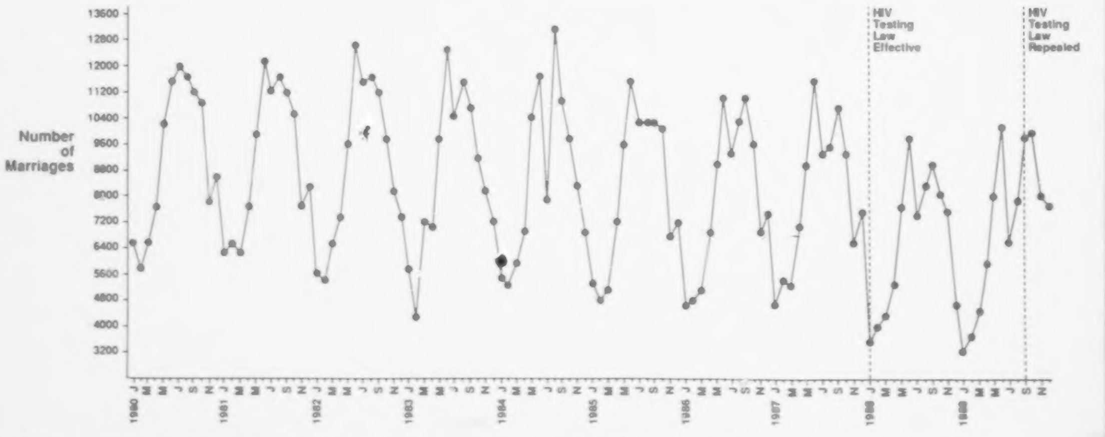

FIGURE 1—Monthly Marriage Time Series for Illinois from January 1980 to December 1989

TABLE 2—Interrupted Time Series Analysis for Marriages in Illinois, as well as Bordering and Comparison States

State	ARIMA Trend12	Testing Law[a]	Model LBQ[b]	Model R^2
Illinois	−241	−1043	26	.941
Indiana	−103	457	19	.857
Wisconsin	−43	324	26	.948
Missouri	−88	317	25	.601
Ohio	−49		22	.945
Minnesota	−52		15	.908
Kansas			19	.826

NOTE: Only statistically significant parameter values are presented. All seven states showed significant serial dependency parameters.

[a] Abrupt, permanent change in monthly marriage time series beginning January 1988.

[b] Chi Square test of ACF residuals at 25 lags. None of the reported values are statistically significant, indicating random residuals.

This monthly drop persisted through August 1989. The model indicates that premarital testing reduced the number of marriages in Illinois by 20,860, about 14 percent over the study period.

Crossover Marriages in Border and Comparison States

Table 2 also presents the results of statistical analyses of marriage time series for the three bordering and three comparison states. An hypothesized crossover effect would be revealed by increases in marriages in the border states and no change in marriages in the comparison states.

Inspection of Table 2 reveals that five of the border and comparison states experienced statistically significant year to year (trend) decline in the frequency of marriage. Kansas was an exception. In that state, monthly marriage frequency cycled around a stable average (level) over the entire study period. All six states showed significant moving average components.

Tests of the crossover hypothesis are provided by the Testing Law parameter. In each of the three border states, analysis revealed an abrupt, permanent, and statistically significant increase in marriages in January 1988. Indiana showed a 10 percent increase, Wisconsin a 9 percent increase, and Missouri an 8 percent increase. Together these increases came to an increase of 1,098 marriages per month, starting in January 1988 and lasting through August 1989. The combined increase in these bordering states almost exactly equals the decrease experienced in Illinois.

Of equal importance for testing the hypothesized crossover effect is the finding that none of the comparison states experienced significant changes in marriages in January 1988. The Testing Law parameter was not significant for any of the comparison states although model prediction (R^2) was as good or better for the comparison states as for the border states.

Rebound of Illinois Marriages

The question of a rebound in the Illinois marriage time series with the repeal of mandatory premarital testing cannot be addressed directly with Box-Jenkins times series analysis without a lengthy, post-law time series. An indirect analysis was made by using the intervention model developed in this study to make predictions for the post-law Illinois experience for the last quarter (October–December) of 1989. This model does not include a rebound parameter. Model predictions were compared with the actual number of marriages reported for Illinois during the final quarter of 1989. A rebound would be observed if the actual number of marriages were greater than those predicted by the interruption model. As shown in Table 3, the no rebound model predicts significantly fewer marriages than actually occurred for each of the three months. The interruption model underpredicted the number of marriages in Illinois by an average of 1,100 per month during the period since the HIV testing requirement was repealed. The rebound in marriages in Illinois was approximately equal to the original drop caused by the testing law.

Discussion

The Illinois premarital testing law produced an abrupt decrease in the number of marriages in Illinois. Analyses revealed a drop in the Illinois marriage time series in January 1988 when the law became effective; the drop occurred immediately and lasted throughout the study period. While severe, about 14 percent, the drop was not as great as the 22.5 percent originally reported.[12] There was no statistical evidence that the drop was temporary or that it diminished or increased over time.

Although there was a significant drop in marriages in Illinois, there was not the predicted absolute drop in marriages. Rather than (a) take the HIV blood test and get married in Illinois, or (b) not get married at all, couples crossed the state line and were married in states bordering Illinois. Analyses reveal a coincident, and approximately equal, abrupt increase in the marriage times series for the states most likely to get marriage crossovers from Illinois—Indiana, Wisconsin and Missouri. These increases did not represent a midwestern trend since January 1988 increases were not found for Ohio, Minnesota, or Kansas.

Analysis suggests that the changes in Illinois marriage patterns produced by the HIV testing requirement were not permanent. With the repeal of the HIV testing law, marriages appear to have rebounded in Illinois.

Cost of HIV Testing

Considerable attention focused on the cost of the HIV test itself as a reason for couples' decisions to cross state lines to marry.[9] The media reported charges of up to $400 a couple for HIV tests.[7,18] However, the financial cost of testing may present too easy an explanation. Turnock and Kelly[12] cite a survey of Illinois hospitals with a mean cost of $30 for each HIV test. For half these hospitals, the charge covered both the ELISA and the Western Blot tests. Even a $30 charge may be high since most couples (99.4 percent) would only require the ELISA test.[5] Mercola[19] reports that the cost of the ELISA test to physicians in Illinois was only $10.

Aside from the financial burden, HIV testing did not increase significantly the difficulty of obtaining a marriage license in Illinois. During this period Illinois required a blood test for syphilis, but so did Wisconsin and Indiana. In Illinois, a marriage license was valid the next day, while a three-to-five-day waiting period was required in the three bordering states (see Table 1).

In order for the cost of the HIV test to have produced the marriage patterns revealed in this report, couples must certainly have incurred logistical and financial expenses greater than the cost of two HIV tests. For example, when an Illinois marriage license cost $15 and required HIV tests, more than 6,000 couples chose instead to be married in Wisconsin. There, a marriage license cost $50 with an additional $10 fee for waiving the five-day waiting period, if the Wisconsin county clerk was willing to grant it.[20]

Fear of AIDS

An alternate explanation for the changes in marriage patterns is fear of an AIDS diagnosis.[12,21] Couples may have

TABLE 3—Study Model Predictions for Post-Premarital Testing Marriage Frequencies in Illinois

1989 Month	No-Rebound Model Predictions	Actual Marriages	Residual (under Prediction)
October	8974	9909	(935)
November	6405	7927	(1522)
December	6771	7615	(844)
Mean	7383	8484	(1100)

NOTE: Standard error for predictions is 664. Residuals indicate that actual marriages were higher than expected from the No-Rebound model.

wanted to avoid learning of their (positive) HIV status. Some evidence supporting this explanation is provided for men by the rise in their seroprevalence rates over the testing period. Turnock and Kelly[12] reported male, marriage license applicants' seroprevalence rates of .85 per 10,000 during the first six months of testing. The Illinois Department of Public Health[22] has released figures indicating male seroprevalence rates of 3.23 per 10,000 over the final 14.3 months of mandatory premarital testing. By contrast, the seroprevalence rate for women remained at about 1.4 per 10,000 during the entire period. Male seroprevalence rates show a four-fold increase from the first six months to the final 14 months of testing suggesting that, at least at the beginning of the testing period, men were avoiding marriage in Illinois because of fear of a positive test result.

Although there is evidence that some men got married outside of Illinois to avoid HIV testing, this explanation does not fit with the large change in marriage patterns any better than does explanation in terms of the cost of the HIV test. One expects that only relatively few men had realistic fears of HIV infection. An explanation that is consistent with the fear of an HIV test result is a generalized fear of AIDS. Perhaps AIDS had become such a specter that the issue itself is to be avoided.[23] Couples may have chosen a way around this unhappy topic by being married in a state that allowed them to without testing. ☐

Jack McKillip, PhD

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I. New York State HIV Seroprevalence Project: Goals, Windows, and Policy Consideration

Lloyd F. Novick, MD, MPH

Introduction

In the summer of 1987, the New York State Department of Health embarked on an extensive effort to determine the prevalence, distribution, and trends of human immunodeficiency virus (HIV) infection in the population. Statewide serosurveillance for antibody to HIV would be conducted by blind testing of available blood samples. Six populations were selected as "windows" on the epidemic: newborns, homeless adolescents, prisoners, and clients of family planning, sexually transmitted disease (STD), and drug treatment clinics. For each of these windows the objectives were to: 1) establish the baseline prevalence and geographic distribution of infection; 2) monitor trends of infection overall and in various populations, including age and racial/ethnic groups; and 3) use the information to target and design preventive activities.

Testing of all newborns in New York State began in November 1987, and all six study windows were operational by June 1988. As of March 1990 blind testing for HIV antibody had been performed on more than 725,000 blood specimens in these six studies. Two years of testing have yielded sufficient data for reliable analysis and projection of trends. The following chapters describe the methods and results of these seroprevalence activities, with emphasis on HIV infection in women of reproductive age, adolescents, and children. HIV seroprevalence rates are correlated with other socioeconomic indicators, including drug-related hospital discharges, the educational level of childbearing women, and the incidence of low birthweights. Information from these surveys has been used to identify high HIV infection rates in small geographic areas and to spur preventive activities statewide.

Impetus for Statewide HIV Serosurveillance

These serosurveillance activities were prompted by two phenomena: the large number of individuals with clinically apparent AIDS in New York State, and the likelihood of a much larger number asymptomatically infected with the virus.

Since the first cases of AIDS were described, New York State has reported more cases than any other state[1] (29,186 by March 31, 1990).[2] Although 84 percent of these cases are residents of New York City, a substantial number (3,564) reside elsewhere in the state. By the same date a total of 18,239 deaths were attributed to AIDS.[2] In New York City, AIDS is the leading cause of death for males 30 to 49 years old, the leading cause of death for females between 20 to 39 years old, and the second leading cause of death for males 20 to 29 years old.[3]

Among children 648 cases of AIDS were reported in New York State as of March 31, 1990.[2] These case reports represent only a fraction of HIV-infected children in New York State, as estimated in the Newborn HIV Seroprevalence Study (see chapter II).

Although case-report data were invaluable in plotting the rise in incidence of AIDS and delineating the associated risk factors, severe deficits in understanding of the epidemic remained. Above all, the current distribution of HIV infection could not be determined from case reports. Most individuals infected with HIV are asymptomatic and not likely to be reported to public health departments, even in states which mandate reporting of HIV infections. (New York State requires reporting of AIDS cases but not HIV infections.) AIDS case reports also lag far behind the prevalence of HIV infection. The median lag time between infection and development of AIDS, resulting in a report to a public health agency, may be 10 years or more.[4]

To determine the actual occurrence of infection, direct and contemporary measures of those infected were needed. The multi-window HIV serosurveillance program could provide such a measure.

A critical need in planning public health interventions was to determine the extent of HIV infection in the entire community and in definable groups and locations. A comprehensive monitoring system, based on a representative, continuing serosurvey, was also needed to track the trends statewide and for age, racial/ethnic, and other groups on a statewide, regional, and community basis.

Approach to Statewide HIV Serosurveillance

The seroprevalence survey method adopted by the New York State Department of Health was blind testing of available blood specimens for antibody to HIV. Representative sampling of households was rejected, although the National Center for Health Statistics and the Centers for Disease Control used this method for a national survey designed to reach 50,000 respondents.[5,6] Seroprevalence analysis of available blood specimens could be implemented faster and more economically than a household survey and would avoid the problems of self-selection and bias associated with community surveys, which require voluntary participation and consent.

Before the Department of Health's blind seroprevalence studies were initiated, extremely limited data were available on the extent of HIV infection in New York State. Seroprevalence information was limited to military recruits and blood donors. Of recruits from New York State 0.35 percent were seropositive, 0.36 percent of males and 0.28 percent of females (Centers for Disease Control: Quarterly Data Report on HIV in Civilian Applicants for Military Service, June 22, 1989). Of 134,144 first-time blood donors seen between July 1987 and September 1988, 0.09 percent were seropositive, 0.12 percent of males and 0.06 percent of females.

The New York State program was based on six major windows, i.e., population groups from whom blood specimens are routinely collected for other purposes. In virtually all cases, sufficient specimen remains for blind assay of antibody to HIV. The six windows are (Table I-1):

Newborns: A dried-blood specimen is routinely collected from every infant born in New York State and screened by the State Department of Health for seven inherited metabolic disorders (see chapter II).

Family planning clients: A number of family planning clinics routinely collect blood specimens for syphilis testing as part of the client's initial medical examination (see chapter VII).

Runaway and homeless adolescents: A shelter for high-risk adolescents in New York City routinely collects blood specimens from individuals referred for medical examination (see chapter IX).

STD clinic clients: Blood is routinely drawn at sexually transmitted disease (STD) clinics for syphilis testing. Twenty clinics outside New York City were recruited in 1987 for the blind HIV study. Results for the fourth quarter of 1989 showed seroprevalence rates of 2.8 percent for 2,792 males and 1.1 percent for 1,318 females. The merged data peaked for clients aged 30–39 (4.3 percent). In the racial/ethnic distribution the rate was higher for Hispanics (3.5 percent) than for Blacks (2.5 percent) or Whites (1.7 percent).[3]

Intravenous drug users (IVDU): Blood is routinely collected from clients being admitted to methadone maintenance programs. Four programs with 35 sites in New York City, including all five boroughs, were recruited for the blind HIV study. Clients entering from June 1988 through December 1989 had very high seroprevalence rates (29.9 percent; males, 31.7 percent; females, 26.6 percent). The rates were high for all age groups, all race/ethnic groups, and all boroughs.[3]

Prisoners: All persons entering the state prison system are processed through central intake facilities, where blood specimens are drawn routinely as part of the medical examination. (For female prisoners, see chapter VI.)

In addition to these six windows, seroprevalance data on two other groups ~~were tapped~~: military recruits and first-time blood donors. Additional data sources have been identified, and the Department of Health is considering incorporation of the data into the survey program.

The Centers for Disease Control developed a similar program in late 1987 to survey HIV seroprevalence for six groups: IVDUs, patients at general hospitals, patients at STD clinics, patients at tuberculosis clinics, patients at clinics serving women of reproductive age, and newborns.[4,5] This Family of Surveys is being conducted in more than 30 areas of the United States. New York is one of the participating states in this surveillance system.

Planning and Policy Issues

Legal Basis—In New York State specimens without identifying information can be used for epidemiologic studies without the need for informed consent. New York State Public Health Law (PHL) Article 24-A regulates research on human subjects and requires informed consent of the subject. However, PHL 2441(2) specifically exempts from the definition of human research "the conduct of biological studies exclusively utilizing tissue or fluids after their removal or withdrawal from a human subject in the course of standard medical practice, or . . . epidemiological investigations." Consequently informed consent of the subjects in the seroprevalence study is not required (memo from Peter J. Millock, General Counsel, New York State Department of Health, to Lloyd Novick, August 24, 1987).

Advisory Committees—An Advisory Committee was assembled on September 17, 1987 by the Department of Health to review the objectives and methods of the proposed seroprevalence surveillance system and the newborn study in particular. The members were Elaine Abrams, MD, Harlem Hospital; Daniel Callahan, PhD, The Hastings Center; Victor DeGruttola, MD, Harvard School of Public Health; Richard Kaslow, MD, National Institute of Allergy and Infectious Diseases; Margarite Pappaioanou, MD, Centers for Disease Control; and Warren Winkelstein, Jr., MD, School of Public Health, University of California, Berkeley.

The Advisory Committee unanimously supported the need for the proposed studies and recommended that data be collected by small area and risk behavior to the extent that anonymity was not compromised. Information on prevalence in areas outside New York City was deemed to be highly important for detecting the extent of spread.

A second special Advisory Committee was convened on March 10, 1988, to review the preliminary results of the newborn serosurvey. In addition to Drs. Abrams and Kaslow, the members were

Table I-1—Windows of the New York State HIV Surveillance Project

Seroprevalence Window	Target Population	Sample	Data Collected	Limitations	Implications
Newborns	Childbearing women 275,000 per annum	All births in NYS 275,000 per annum	Age of mother, race/ethnicity of infant, zip code of hospital	Applicability to reproductive age women	a) Childbearing women b) Women of reproductive age c) Pediatric HIV infection
Family Planning Clinics	Women receiving family planning	41 sentinel family planning clinics 13,000 per annum	Age, race/ethnicity, zip code of family planning center	Sentinels are not representative of all women in family planning	a) Women in family planning b) Women of reproductive age
Homeless Adolescents	Homeless and runaway youth	NYC adolescent facility 12,000 per annum	Age, race/ethnicity, gender, risk	Not representative of the adolescent population	a) Document infection in adolescent groups b) Sentinel population to detect changes over time c) Relationship of risk to prevalence of infection
STD Clinics	Clients at county STD clinics	20 STD clinic sites outside of NYC 5,000 per annum	Age, race/ethnicity, sex, zip code of clinic, syphilis	Representative of only those who go to STD clinics	a) Men and women in STD clinics b) Association of HIV-ab and syphilis c) Spread in women, and, by inference, via heterosexual transmission
IVDU Clinics	All IVDU clients of treatment (methadone)	35 methadone maintenance treatment programs & detox programs in NYC	Age, race/ethnicity sex, treatment modality, drugs injected, sexual orientation, needle sharing, treatment (drug) Hx of HIV-related illness	Clients are in methadone-treatment programs (main drug of abuse = heroin, not cocaine). Not representative of addicts not in treatment	a) IVDU newly entering treatment b) IVDU already in treatment
Prisons	NYS prisoners	Sample of inmates entering at the four intake facilities 500 per annum	Prison site, residence (county), sex, age, race, risk group	Applicability to other populations	a) NYS prisoners b) High-risk population (IVDU), minorities c) Fertility information available

Timothy Dondero, Jr., MD, Centers for Disease Control; Myron Essex, DVM, PhD, Harvard School of Public Health; Harvey Fineberg, MD, Harvard School of Public Health; Margaret Heagarty, MD, Harlem Hospital; Keith Krasinski, MD, Bellevue Hospital; Margaret Oxtoby, MD, Centers for Disease Control; Peter Selwyn, MD, Montefiore Medical Center, and Issac Weisfuse, MD, New York City Department of Health.

This Advisory Committee strongly recommended continuance of the newborn serosurvey, particularly for the purpose of monitoring the epidemic.

Laboratory Facilities—A capability to perform the necessary volume of HIV screening and confirmatory tests was developed at the Department's Wadsworth Center for Laboratories and Research. Automation of the testing process, including employment of robotics, facilitated the implementation of a large seroprevalence study. The Wadsworth Center is now performing nearly 400,000 serologic tests for antibody to HIV annually.

Confidentiality—All laboratory HIV determinations are entered into a computerized data-management system. Names of individuals are not collected, and no identifying information is entered into the data system. Demographic data are collected for all studies (age, sex, race/ethnicity, and zip code), and risk factors are collected for some windows (homeless adolescents, drug treatment clients, and prisoners). To prevent potential identification of any individual from demographic information on a single specimen or a small number of specimens, data are reported only on subgroups that include at least six individuals.

Public Information—Extensive public information was provided about the nature and rationale for this effort before inception of the studies. Information on the newborn study was also provided to hospitals and hospital-based pediatricians. Some public concern was initially expressed about the blind nature of these studies and the inability to notify individuals of their results. To allay this concern, voluntary HIV testing and counseling was offered to individuals in most of the participating facilities.

Ethics—Ethical questions have been raised regarding blind surveys, because positive results indicating infection cannot be linked back to the individual for medical follow-up. Under New York State law individuals cannot be notified of test results from nonidentified (blind) samples obtained without consent. The solution taken by the Department of Health was to offer voluntary testing, wherever feasible, to individuals at the same time as the blind testing. Review of the issues raised by blind testing in a number of reports have emphasized the epidemiologic importance of this type of study and its consistency with ethical practice.[7–9]

Data Limitations and Uses—The limitations of data from blind seroprevalence studies were recognized at the outset. The chief limitation is that the groups available for study are determined by the availability of blood specimens. Some population groups, such as adult males, may be difficult to assess, and the study group is not representative of the entire population. There are also limits on the type of information that can be collected, particularly detailed risk histories.

Despite these limitations the serosurvey results are providing critical data on HIV prevalence in the subgroups tested and on changes over time. The newborn study in particular may have wide applicability to the general population of women of reproductive age in the community, since virtually all childbearing women are participants. Issues of completeness and general applicability of the data are dealt with in later chapters.

Collectively these seroprevalence activities are providing invaluable information on the extent of the HIV epidemic in New York State. Information from these studies has directly influenced the development of preventive actions and public policy.

Summary

In the summer of 1987, the New York State Department of Health embarked on an extensive effort to determine the prevalence of trends of HIV infection in the state. Serosurveillance for antibody to HIV was undertaken by blind testing of available blood samples. Six populations were selected as windows on the epidemic: newborns, homeless adolescents, prisoners, and clients of family planning, STD, and drug treatment clinics. Testing of all newborns in New York State began in December 1987, and all six study windows were operational by June 1988. As of March 1990 blind testing for HIV antibody had been performed on more than 725,000 blood specimens in these six studies. Collectively these seroprevalence activities are providing invaluable information and impetus to preventive actions and public policy. ☐

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II. Newborn Seroprevalence Study: Methods and Results

Lloyd F. Novick, MD, MPH
Donna M. Glebatis, MS
Rachel L. Stricof, MPH
Patricia A. MacCubbin, MS
Lawrence Lessner, PhD, MPH
Donald S. Berns, PhD

Introduction

The New York State Department of Health's ongoing study of HIV infection in childbearing women is based on testing of blood specimens obtained from all infants born in this state for seven inherited metabolic disorders. The specimens are collected and initially tested by the Department's Newborn Screening Program. The specimens are then sent to the HIV Seroprevalence Laboratory for blind HIV testing. This paper describes the results for specimens collected during the 28-month period from November 30, 1987 through March 31, 1990.

Since maternal IgG antibodies cross the placenta, the presence of antibodies in a newborn is indicative of infection in the mother. These antibodies do not necessarily indicate that the child is infected. However, an estimated 30–50 percent of seropositive newborns are actually infected.[1-3]

Methods

Testing Protocol

The protocol for HIV testing of newborn blood specimens has been described in detail by Novick, *et al.*[4] Briefly, a dried-blood specimen is obtained from each infant along with certain demographic information, including the mother's age, zip code and county of residence, the newborn's race/ethnicity, and the month and year of birth. Hospitals are informed of the need for accurate, complete reporting and are contacted, if necessary, to obtain missing information. For the HIV study the demographic information—but not identifying information—is entered into a separate, unlinked data base, where each record is identified by a unique, randomly generated number. The same number is assigned to a portion of the blood specimen. All metabolic disorder screening tests are completed before specimens are submitted for HIV testing.

Specimens for which insufficient or unsuitable blood is left after the metabolic disorder screening are excluded from the blinded HIV seroprevalence study. The number of these specimens is carefully tracked. Repeat specimens cannot be obtained at this stage because the blinding process makes it impossible to link the identity of the infant to the specimen. To ensure that only one specimen representing each childbearing woman is tested, the following are also excluded: all specimens except one from a multiple birth event and all repeat specimens requested for follow-up by the Newborn Screening Program (Figure 1).

Repeat Live Births during Study Period

The likelihood that some women would present with a live birth on more than one occasion during the study period was estimated using Vital Records birth files for 1988 and 1989. Although the Vital Records and HIV newborn data are not directly comparable, this method provides a rough estimate of the percentage of women in the HIV newborn study who had two or more separate live birth deliveries during the study period.

Overall, 3.7 percent of the live-birth records indicated a previous live birth within the same two-year period, 1988–89. However, 29 percent of the New York City records and 11 percent of the remaining records were excluded from this analysis because of incomplete or inaccurate information on the date of the last live birth.

Definition of Test Results

Since the initiation of the Newborn Seroprevalence Study, three sets of crite-

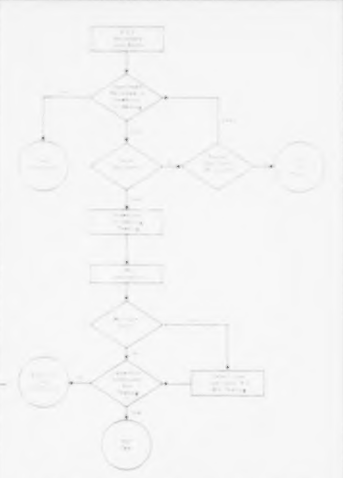

FIGURE 1—New York State Newborn Screening/Newborn HIV Seroprevalence Testing Process

ria have been used to interpret the Western blot results. Reactives were defined as follows:

- Criteria I (November 30, 1987–March 31, 1988): All three of the following: Envelope (gp 41 or gp 120/gp 160), core (p24), and polymerase (p31, p51, or p66).
- Criteria II (April 1, 1988–June 26, 1989): Envelope and either core or polymerase, as defined above.
- Criteria III (June 27, 1989 to present): Any two of the following: p24, gp41, and (gp120 and/or gp160).[5]

Equivocals were defined as virus-specific bands other than the set described above. Nonreactives displayed no virus-specific bands.

Since the individual Western blot band results are currently not available prior to August 1988, the effect of the change from Criteria I to Criteria II could not be examined. However, the change from Criteria II to Criteria III was examined by applying Criteria II to the specimens tested June 27, 1989 to present and comparing the number and percent HIV positive to that found under Criteria III. Under Criteria II there were 1,918 HIV positives with a seroprevalence rate of 0.661 percent. This was nearly identical to that found under Criteria III with 1,921 HIV positives and a seroprevalence rate of 0.662 percent.

Statistical Analysis

Seroprevalence rates were calculated based on the number of Western blot reactives and total specimens tested, excluding equivocals. We could not obtain a second sample for reevaluation of equivocals because of the blinded nature of the study.

Age/race-specific HIV seroprevalence rates were calculated for five racial/ethnic subgroups (White, Black, Hispanic, Other, and Unknown), and six maternal age categories (<20 years, 20–24 years, 25–29 years, 30–34 years, 35+ years and Unknown).

Geographic rates were calculated for New York City total, each of the five New York City boroughs, New York State exclusive of New York City, and four regions outside New York City, which were selected on the basis of urbanization, physical proximity to New York City, and cumulative incidence of AIDS to date. The four regions were New York City Suburban (Nassau, Suffolk, Westchester, and Rockland counties), Mid-Hudson Valley (Dutchess, Putnam, Orange, Sullivan, and Ulster counties), Upstate Urban (Albany, Schenectady, Onondaga, Monroe, and Erie counties), and Upstate Rural (all other counties in the state).

For New York City the HIV seroprevalence rates were also calculated for each zip code area and mapped by quartiles. Zip code areas with fewer than 100 births were eliminated from this mapping. For New York State exclusive of New York City, due to the large number of zip code areas and the small number of births within each area, only the rates for areas with two or more HIV-positive cases and rates twice that of the overall rate were examined, and the rates were mapped by county.

Trend analysis was conducted on the monthly HIV seroprevalence rates for births occurring in the 27-month period, December 1987 through February 1990. Births occurring in November 1987 were excluded from this analysis because the specimens received beginning November 30 covered two-thirds of the November newborns. Plots of the monthly rates were analyzed by simple regression techniques[6–8] with SAS software.[9] A coefficient of determination (R^2) was calculated to determine the proportion of the variation in rates explained by the independent variable, time (month of birth). The 95 percent confidence intervals for the predicted value were also calculated.

This procedure was performed for all of New York State, all of New York City, all of New York State exclusive of New York City, each of the five New York City boroughs, and each of the four other regions described above. The data for all of New York State, all of New York City, and all of New York State exclusive of New York City were further analyzed for three maternal age groups (<20, 20–29, 30+) and for three race/ethnicity groups (White, Black, and Hispanic).

Results

During the study period November 30, 1987 through March 31, 1990, a total of 653,328 blood specimens were analyzed for HIV serologic status, of which 211 equivocals (0.03 percent) were eliminated from the analysis. Of the remaining 653,117 specimens, 10,125 were from women residing outside of New York State, and 2,408 were from women whose residence was unknown. These specimens were also eliminated from the analysis. For the remaining 640,584 specimens, the overall seroprevalence rate was 0.66 percent: 1.24 percent in New York City and 0.17 percent in the rest of the State.

New York City

The HIV seroprevalence rates for New York City and each borough are shown in Table I-1. The rates were highest in the Bronx, followed by Manhattan and then Brooklyn. The rates for Queens and Staten Island were more than twice the rates for the two New York State regions adjacent to New York City, New York City Suburban, and Mid-Hudson Valley.

Of the 170 zip code areas in New York City with 100 or more births, 31 (18.2 percent) had HIV seroprevalence rates in excess of 2 percent and 13 (7.6 percent) had rates in excess of 3 percent. The distribution by quartiles is shown in Figure 2.

The vast majority of HIV seropositives in New York City (89.3 percent) were Black or Hispanic, although only 60.8 percent of newborns were Black or Hispanic.

The highest rates were detected among Black and Hispanic newborns. These rates increased with maternal age for Blacks and Hispanics but not for Whites or Other (Table II-2). The rates for both Black and Hispanic infants peaked in the 30–34 year maternal age group.

Age-specific seroprevalence rates are plotted in Figure 3. The sharpest rise is seen in New York City for infants born to mothers between ages 14 and 24. These rates increased ninefold, from 0.16 percent (1 in 624) for 14-year-olds to 1.41 percent (1 in 71) for 24-year-olds.

TABLE II-1—Percent of HIV Positive Newborns by Region for New York State, November 30, 1987–March 31, 1990

Region	Number Positive	Total Tested	Percent Positive
New York State Exclusive of New York City	601	346,522	0.17
NYC Suburban	329	120,422	0.27
Mid-Hudson Valley	71	29,450	0.24
Upstate Urban	119	88,088	0.14
Upstate Rural	82	108,562	0.08
New York City	3,650	294,062	1.24
Manhattan	799	50,364	1.59
Bronx	998	58,003	1.72
Brooklyn	1,352	104,613	1.29
Queens	424	67,474	0.63
Staten Island	77	13,608	0.57

Women with White seropositive infants were slightly younger (mean 28.6 years) than women with White seronegative infants (29.2 years). Women with Black seropositive infants were significantly older than women with Black seronegative infants (28.2 versus 26.0 years). Women with Hispanic seropositive infants were also significantly older (27.2 versus 25.7 years).

New York State Exclusive of New York City

Seroprevalence rates for New York State and its four regions exclusive of New York City are shown in Table II-1. Rates for two regions adjacent to New York City (New York City Suburban and Mid-Hudson Valley) were only about one-fifth the rate for New York City but twice or three times the rates for the other regions.

HIV seroprevalence was concentrated in certain areas of the state (Figure 4). In New York State exclusive of New York City, 64 zip code areas had two or more seropositives and a HIV seroprevalence rate more than double the average for all areas outside New York City. Of the 64 zip code areas, 50 had rates of 0.5 percent or higher. The 64 zip code areas contained 65 percent of the HIV seropositives but only 16 percent of the newborns tested outside of New York City.

In New York State exclusive of New York City, HIV seroprevalence was highest for Black newborns followed by Hispanics (Table II-3). The majority of HIV seropositives (69.4 percent) were either Black or Hispanic, although only 14.0 percent of the newborns tested were Black or Hispanic. For Black newborns the HIV seroprevalence rate increased with maternal age, peaking in the 35+ age group. For Hispanic newborns the rate increased only in this maternal age group. For White newborns no significant variation with maternal age was found.

As seen in New York City, mothers of White seropositive infants were on average slightly younger than those with White seronegative infants (26.5 years vs 27.7 years); however, the difference is significant in women in New York State exclusive of New York City. Women with Black seropositive infants were significantly older than those with Black seronegative infants (27.7 years vs 24.6 years). Mothers of Hispanic newborns were 1.1 years older (26.5 years vs 25.4 years).

Trend Analysis

When HIV seroprevalence rates were plotted by month of birth for New York City and for New York State exclusive of New York City, with 95 percent confidence intervals, neither slope was significantly different from zero. No significant increase or decrease was found for any borough of New York City, any region outside of New York City, or any three racial/ethnic subgroups in New York State exclusive of New York City.

Only two age groups showed evidence of a statistically significant trend for the 27-month period. For New York City women less than 20 years old (Figure 5A), the rate declined slightly ($y = 0.00785 - 0.000132 \times$, $p < .01$). The standard error of the slope was 0.0000486. However, the coefficient of determination (R^2) was only .23, explaining only a small proportion of the variation in the rate.

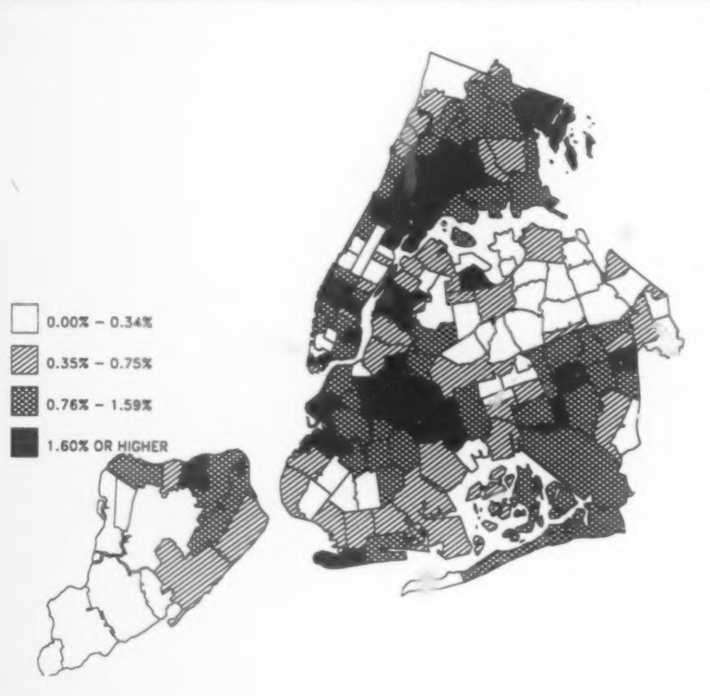

FIGURE 2—HIV Seroprevalence in New York City by Zip Code, November 30, 1987–March 31, 1990

TABLE II-2—Distribution of HIV Positive Newborns by Maternal Age and Race/Ethnicity for New York City, November 30, 1987–March 31, 1990

Race/Ethnicity	HIV Positive by Maternal Age Group (years)						
	<20	20–24	25–29	30–34	35+	Unknown	Total
White							
HIV Positive							
Percent	0.22	0.42	0.38	0.29	0.28	0.74	0.34
Number	6	58	102	74	41	3	284
Tested	2,716	13,935	27,158	25,715	14,403	408	84,335
Black							
HIV Positive							
Percent	0.68	1.68	2.59	3.08	2.98	3.10	2.19
Number	95	429	670	547	268	19	2,028
Tested	14,002	25,512	25,831	17,780	8,985	612	92,722
Hispanic							
HIV Positive							
Percent	0.57	1.26	1.75	1.95	1.42	2.35	1.43
Number	74	333	425	287	101	12	1,232
Tested	12,909	26,519	24,232	14,720	7,113	511	86,004
Other							
HIV Positive							
Percent	0.37	0.43	0.28	0.14	0.23	0.00	0.27
Number	4	23	29	11	8	0	75
Tested	1,095	5,293	10,222	8,037	3,408	147	28,202
Unknown							
HIV Positive							
Percent	0.51	1.10	1.52	0.69	1.76	1.01	1.11
Number	1	5	9	3	4	9	31
Tested	195	454	594	435	227	894	2,799
Total							
HIV Positive							
Percent	0.58	1.18	1.40	1.38	1.24	1.67	1.24
Number	180	848	1,235	922	422	43	3,650
Tested	30,917	71,713	88,037	66,687	34,136	2,572	294,062

For women in New York State exclusive of New York City aged 20–29 (Figure 5B), the rate increased slightly ($y = 0.00149 + 0.00003 \times, p < .03$). The standard error of the slope was 0.0000128. R^2 for this group was extremely small (.17).

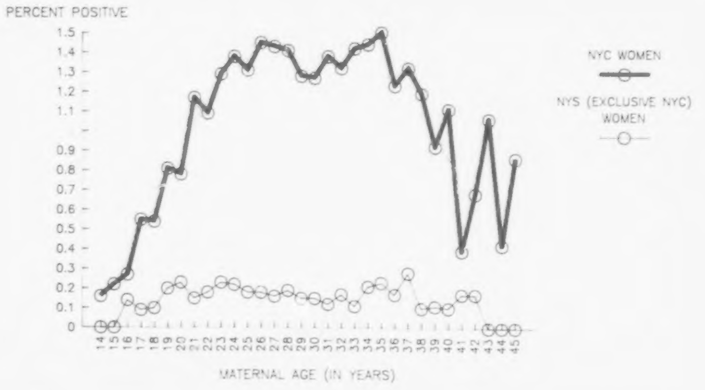

FIGURE 3—HIV Seroprevalence in NYS by Single Year of Maternal Age (November 30, 1987–March 31, 1990)

Discussion

Before this study was implemented, extremely limited information was available on the prevalence of HIV infection in childbearing women. Two recent studies in New York City hospitals revealed an HIV seroprevalence rate of 2.4 to 2.5 percent in infant cord blood specimens.[10,11] A recent Massachusetts Department of Health study found an overall rate of 0.21 percent in anonymously tested newborn blood samples, ranging from 0.1 percent in suburban and rural hospitals to 0.8 percent in inner-city hospitals.[12]

In this study, we found an overall HIV seroprevalence rate of 0.66 percent for childbearing women. The magnitude of this health problem was most striking in New York City where in some zip code areas as many as one out of 22 childbearing women were found to be HIV-infected. For New York State exclusive of New York City the lower overall seroprevalence rate (0.17 percent) is deceptive, because zip code analysis revealed small areas with rates approaching those in New York City. This small-area information is critical for assessment of large geographic areas, where apparently low seroprevalence may mask communities with rates of serious public health concern.

Although 87 percent of the seropositive newborns in New York State were Black or Hispanic, the seroprevalence rate of 0.12 percent for Whites and for other racial/ethnic groups confirms a significant incursion of HIV infection in those populations as well.

The finding of an increased infection rate with increasing maternal age has implications for the design of preventive activities. An especially sharp rise was seen for adolescents and young women in New York City, from 1 in 624 at age 14 to 1 in 71 at age 24 (Figure 3). Preventive efforts must begin with children and younger adolescents to control this rapid spread.

Although racial/ethnic and maternal-age patterns are crucial for projections and resource allocation, the intersection of these patterns offers deeper insight into the environmental factors that determine prevalence. Identifying such cohorts will help to define the relevant exposures, such as patterns of drug use, and thus the targeting of preventive measures. Individuals in younger cohorts with different exposure risks may exhibit substantially different patterns of HIV infection in the future. Those changes will affect not only projections of the epidemic, but also the measures needed to reduce transmission of the virus.

Certain biases inherent in this study may have an impact on the observed rates. Possible bias resulting from newborns not

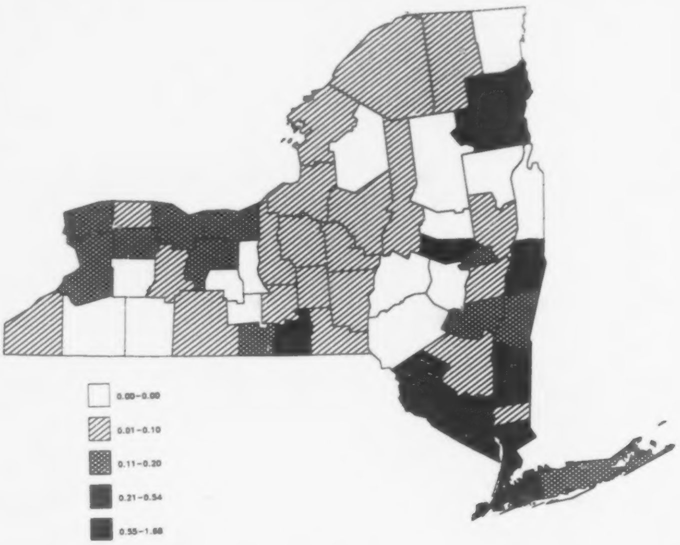

FIGURE 4—HIV Seroprevalence in NYS exclusive of NYC by County (November 30, 1987–March 31, 1990)

TABLE II-3—Distribution of HIV Positive Newborns by Maternal Age and Race/Ethnicity for New York State Exclusive of New York City, November 30, 1987–March 31, 1990

	HIV Positive by Maternal Age Group (years)						
Race/Ethnicity	<20	20–24	25–29	30–34	35+	Un-known	Total
White							
HIV Positive							
Percent	0.05	0.09	0.06	0.04	0.03	0.09	0.06
Number	9	53	57	35	10	1	165
Tested	18,943	59,858	101,038	78,112	29,328	1,095	288,374
Black							
HIV Positive							
Percent	0.26	0.67	1.30	1.60	1.72	1.52	0.95
Number	20	75	120	82	39	2	338
Tested	7,660	11,125	9,245	5,132	2,269	132	35,563
Hispanic							
HIV Positive							
Percent	0.51	0.59	0.55	0.55	1.32	0.00	0.61
Number	10	25	20	11	13	0	79
Tested	1,966	4,243	3,639	1,986	982	51	12,867
Other							
HIV Positive							
Percent	0.00	0.45	0.18	0.13	0.09	0.00	0.19
Number	0	7	5	3	1	0	16
Tested	561	1,563	2,758	2,285	1,058	44	8,269
Unknown							
HIV Positive							
Percent	0.00	0.48	0.00	0.00	0.79	0.20	0.21
Number	0	1	0	0	1	1	3
Tested	56	207	316	246	127	497	1,449
Total							
HIV Positive							
Percent	0.13	0.21	0.17	0.15	0.19	0.22	0.17
Number	39	161	202	131	64	4	601
Tested	29,186	76,996	116,996	87,761	33,764	1,819	346,522

tested is discussed in a later chapter.[13] Another bias could result from double-counting the estimated 3.7 percent of women in the study who may have had two or more separate live birth deliveries within the study period.

Regardless of these biases, data from the seroprevalence study are invaluable in developing estimates of the number of infected women of reproductive age and the number of HIV-infected children. If we assume that 30 percent of seropositive newborns will prove to be infected, 1,275 children with HIV infection were born in New York State during the study period. Yet, only 648 cases of pediatric AIDS had been reported as of March 1990. This gap, attributable mainly to infected but asymptomatic children or those with health problems but no HIV disorder diagnosis, is highly relevant to our understanding of pediatric HIV infection and the need for health care resources.

The few trends found in this study reflect changes in seroprevalence over a brief period of only 27 months. Longer monitoring may be necessary to detect changes in this epidemic. The seroprevalence rates are for a population of childbearing women who have already been infected, plus those infected during the study period, minus those who leave the population by outmigration, death, or aging out of the study. A stable seroprevalence in this limited study may, therefore, be compatible with an actual rising or falling incidence of new infection. The observed trends should also be viewed with caution, because two significant results in 30 separate analyses may be the result of chance alone.

Little information on HIV seroprevalence in newborns is available for the period before late 1987. A rapid increase in HIV infection in childbearing women is assumed to have occurred at some point between 1980 and the present. At the end of 1983, only 174 female AIDS cases had been reported to the New York State AIDS Registry.[14] By the end of March 1990, that number had increased 24-fold to 4,194.[15] The lack of observed trend in the 27-month study period and the increase in infection manifested by AIDS in childbearing women prior to this point suggest a decrease in the rate of increase of the epidemic of HIV infection in childbearing women. However, progress of the epidemic is not likely to be uniform throughout the state with certain lower-prevalence areas possibly exhibiting upward trends in the future.

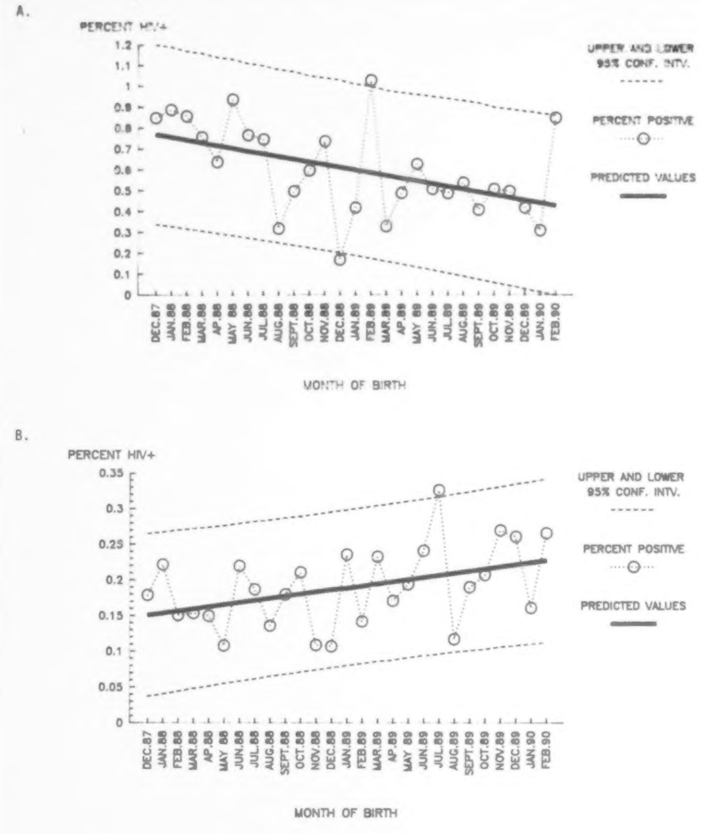

FIGURE 5—Percentage of HIV-Positive Newborns in NYS by Month of Birth for Two Groups with Statistically Significant Trends (December 1987–February 1990)

A—Women less than 20 years old, NYC
B—Women 20–29 years old, NYS exclusive of NYC

Newborn infants of adolescent mothers in New York City showed a decreasing trend in HIV seroprevalence during the 27-month period. This group is particularly critical to the monitoring process as newborns of teenaged mothers have the lowest seroprevalence rates but the greatest rate of increase with maternal age. Seroprevalence among this group may be a closer reflection of recent incidence, since the time interval between HIV infection and childbearing is most likely shorter than that for the older women.

Possible saturation of susceptible individuals in certain geographic areas may also offer an explanation for the lack of observed trends. If so, our efforts toward prevention must be redoubled in these areas, as well as communitywide, to prevent transmission from this pool to others entering various risk groups in the community.

Summary

For the 28-month period, November 30, 1987 through March 31, 1990, 653,117 blood specimens obtained on all newborn infants in New York State for detection of metabolic disorders were also analyzed for HIV serologic status. The overall seroprevalence rate was 0.66 percent: 1.24 percent in New York City and 0.17 percent in New York State exclusive of New York City. Rates of seropositivity were highest in the Bronx (1.72 percent) and Manhattan (1.59 percent). Outside of New York City, HIV seropositivity was concentrated in certain areas. Sixty-four zip codes with two or more seropositives and an HIV seroprevalence rate twice the average outside of New York City contained 65 percent of the HIV seropositives but only 16 percent of the newborns tested. Newborn seropositivity increased with maternal age. In New York City, the seroprevalence rates increased from 0.16 percent (1 in 624) for 14-year-olds to 1.41 percent (1 in 71) for 24-year-olds, a ninefold rise. This survey has provided the impetus for a number of preventive initiatives. ☐

Acknowledgments

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III. Comparison of Newborn Screening Records and Birth Certificates to Estimate Bias in Newborn HIV Serosurveys

Kenneth A. Pass, PhD
Lewis M. Schedlbauer, BA
Patricia A. MacCubbin, MS
Donna M. Glebatis, MS

Introduction

For the New York State Newborn HIV Seroprevalence Study of the magnitude of the HIV epidemic among childbearing women, the completeness and validity of the data as representative of the target population are of prime concern. Missed opportunities for HIV antibody testing fall into three categories:

1. The specimen was not received by the Newborn Screening Program;
2. The initial specimen was invalid or unsuitable for testing and a repeat specimen could not be obtained; or
3. The specimen was insufficient or unsuitable for HIV testing after the mandatory testing was completed.

To assess the potential bias from each of these sources, we determined the number and percent of newborns not tested for HIV antibody within each category and compared the distributions of birth site, residence, race or ethnicity, and birthweight for those not tested and those tested.

Methods

To determine the percent of newborns for which an initial specimen was never received (category 1), live-birth records for 1,000 New York State residents were selected from all 1988 vital records by use of a random sample adequate to yield 95 percent confidence intervals with 1 percent error. These records were computer-matched to newborn screening records; for unmatched records, telephone calls were made to the hospital of birth. The number and percent of live-birth records for which no newborn screening record was available were calculated.

For newborn screening specimens initially invalid with no repeat follow-up specimens and for specimens insufficient or unsuitable for HIV testing, the distributions of birth site, residence, race or ethnicity, and low birthweight (less than 2,500 grams) were compared to the distributions for HIV-tested specimens.

To estimate the minimum impact of the bias, the untested specimens in category 2 (invalid-no repeat) and category 3 (unsuitable/insufficient) were grouped by race or ethnicity and region of birth. The two regions were New York City and New York State exclusive of New York City. The observed HIV seropositivity rate for each race/region subgroup was applied to the number of untested specimens in that subgroup. The overall seropositivity rates were then recalculated with the addition of these results.

To estimate the maximum potential impact of bias from all untested newborn, including those in category 1 (never received), an alternative assumption was made that all those untested were HIV-seropositive. The overall rates were then recalculated.

Results

Specimens Not Tested for HIV Antibody

From the 1,000 randomly selected live-birth records, specimens submitted for newborn screening were matched for 985 newborns. For the remaining 15 birth records two specimens were lost in transit, six were never submitted, and seven represented infants born outside New York State. The parents of these seven infants reside in New York State, but no newborn screening specimen would be expected for an out-of-state birth. Thus eight specimens from an expected 993 (0.82 percent) were not received for newborn screening, but 15 of 1,000 specimens

TABLE III-1—Racial/Ethnic Distribution of Newborns in New York State Tested and Not Tested* for HIV by Region of Birth

		Racial/Ethnic Group (%)				
Region and Group	Number	White	Black	Hispanic	Asian	Other/ Unknown
NYC						
Tested	146,000	29	32	28	5	5
Not Tested	2,506	24	38	30	3	5
NYS Exclusive of NYC						
Tested	123,883	84	10	3	1	2
Not Tested	3,392	80	13	4	1	2
Unknown						
Tested	5,747	68	9	7	5	11
Not Tested	393	21	2	1	1	75

*Invalid and no repeat specimen available, or specimen insufficient or unsuitable.
NYC = New York City
NYS = New York State

(1.5 percent) for infants born to New York State residents were not received for HIV testing.

Of the 281,921 initial specimens received in 1988, a total of 275,630 were tested for HIV, while 1,558 (0.5 percent) were invalid and no repeat specimen could be obtained. An additional 4,733 (1.7 percent) specimens had an insufficient quantity left over after mandatory testing or were unsuitable for HIV testing. If the total number of specimens received was increased by 1.5 percent to reflect the estimated number of specimens never received, the expected total of HIV specimens would be 286,150. Of these, 275,630 were actually tested, leaving an estimated 10,520 newborns (3.7 percent) not tested.

Marked differences were found among hospitals in the percentage of newborns not tested (data not shown). Thus the seroprevalence rates were impacted by factors other than the biological transfer from mother to child.

Comparison of HIV-Tested and Not Tested Newborns

For category 2 (invalid-no repeat) and category 3 (specimen insufficient or unsuitable) considered together, 54 percent of untested and 45 percent of tested newborns were from New York City. More untested than tested newborns were associated with an unknown or out-of-state residence (6 percent and 2 percent, respectively).

The racial/ethnic distributions of HIV-tested and not tested newborns by region are presented in Table III-1. In both New York City and New York State exclusive of New York City the untested groups had a higher percentage of Blacks.

Category 2 (invalid specimens with no repeat specimens obtained) had a disproportionate share of specimens at lower birthweight (13 percent, compared to only 9 percent of all specimens received). Birthweight information was not available for 6 percent of the invalid-no repeat group and for only 2 percent of all specimens.

Estimation of Potential Bias

If the assumption is made that untested specimens in categories 2 and 3 had HIV seroprevalence rates equal to those observed for their race/region subgroup, the number of additional HIV-seropositive infants expected would be 54 (Table III-2). Including these untested newborns would increase the overall seroprevalence rate from 0.642 (1,770/275,630) to 0.647 (1,824/281,921).

The estimated number of all newborns not tested for HIV (including category 1) in 1988 is 10,520. Assuming all 10,520 to be seropositive and including them in the calculation would increase the overall seroprevalence rate nearly sevenfold, from 0.64 percent to 4.29 percent.

Discussion

Since our newborn screening program is designed to test all infants born in New York State, use of the dried blood spots for HIV seroprevalence testing ensures that there is no predetermined bias. However, bias may exist due to biologic factors, such as the mother's health or exacerbating medical conditions of the child. This bias may be particularly important in areas with low seroprevalence rates, where a systematic error could distort the calculations.

The results reported here indicate that the seroprevalence study was able to determine the HIV serostatus of all but 3.7 percent of the childbearing women in the state in 1988. Although the proportion not tested is relatively small, it could still be an

TABLE III-2—Estimated Number of HIV+ Newborn among Those Not Tested in Categories Two* and Three**

Region/Race/ Ethnic	Number Not Tested	Percent HIV+	Estimated No. of HIV+
NYC			
White	814	0.34	2.8
Black	1,289	2.17	28.0
Hispanic	1,018	1.43	14.6
Asian	102	0.11	0.1
Other/Unknown	170	0.62	1.1
NYS Exclusive of NYC			
White	2,005	0.06	1.2
Black	326	0.87	2.8
Hispanic	100	0.53	0.5
Asian	25	0.03	0.0
Other/Unknown	50	0.30	0.2
Unknown			
White	82	0.19	0.2
Black	9	4.46	0.4
Hispanic	5	2.22	0.1
Asian	3	0.00	0.0
Other/Unknown	293	0.83	2.4
Total	6,291		54.4

*Initial specimens invalid with no repeat specimens.
**Insufficient or unsuitable specimen for HIV testing.
NYC = New York City
NYS = New York State

important factor if the characteristics of the untested newborns are similar to those population subgroups with high HIV seroprevalence.

Our study indicated that bias may exist from the untested newborns because their racial/ethnic, regional, and birthweight distributions differed from those tested. If these untested newborns were all seropositive, the seroprevalence rate would be increased nearly sevenfold. This is the maximum bias in these data. The bias of 0.005, calculated by applying the existing seroprevalence rates to specimens in categories 2 and 3, is doubtless unrealistically low. The true bias lies somewhere in between.

Initiation of testing newborns for HIV status has focused additional attention on the quality of specimens and completeness of data submitted to the Newborn Screening Program. Several improvements have been observed. At the beginning of the seroprevalence study 12 percent of the specimens submitted were missing key information such as mother's age, county of residence, and ethnicity. Hospitals were called to secure the information. Within four months less than 3 percent of the specimens were submitted with information missing. Similarly, the quality of blood specimens increased due to focused efforts of laboratory staff in dealing with problem facilities.

Summary

Implicit in the New York State Newborn HIV Seroprevalence Study is the assumption that newborns of all New York State residents are tested for human immunodeficiency virus (HIV) antibodies. We examined this assumption by describing that part of the 1988 New York newborn population not tested in the HIV seroprevalence study and assessing any bias contributed by this group. Of the expected total HIV specimens 1.5 percent were never received by the Newborn Screening Program, 0.5 percent were invalid specimens for which no repeat specimen could be obtained, and 1.7 percent were unsuitable or of insufficient quantity to be tested for HIV antibody. Thus 96.3 percent of all 1988 New York newborns were tested for HIV antibody. Black infants from New York City and low-birthweight infants were represented disproportionately among those not tested. Assignment of all untested newborn to HIV-positive status increased the seroprevalence rate 17 percent (0.64 percent to 0.75 percent). ☐

Acknowledgments

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IV. Geographic Distribution of Newborn HIV Seroprevalence in Relation to Four Sociodemographic Variables

Dale L. Morse, MD, MS
Lawrence Lessner, PhD, MPH
Michael G. Medvesky, MPH
Donna M. Glebatis, MS
Lloyd F. Novick, MD, MPH

Introduction

Since the AIDS epidemic was first recognized, there has been geographic clustering of AIDS cases within the United States and New York State. Most of this clustering is related to pools of HIV-infected individuals and to risk behavior activities which occur most frequently in metropolitan areas and states with large populations at risk. Witness the high rates of AIDS cases among homosexual and bisexual men in New York City and San Francisco and among intravenous drug users in urban areas of New York, New Jersey, and Florida.[1,2] In an earlier study in New York City, zip code areas with rates of drug-related hospital discharges in the top quartile had a newborn HIV seroprevalence of 2.2 percent versus 0.8 percent in all other zip code areas.[3]

Identification of geographic clusters of cases can be useful in focusing preventive efforts and allocation of health care resources. HIV/AIDS cases are also likely to be closely clustered in association with other sociodemographic variables. Knowledge of such relationships could be useful for targeting and design of prevention services which would differ between high-risk areas and areas where HIV/AIDS has not yet surfaced.

This report summarizes data on the geographic distribution of HIV seroprevalence among New York State newborns and its association with four sociodemographic variables: birthweight, maternal education, race/ethnicity, and drug abuse leading to hospitalization. Although such associations do not indicate causation, they provide insights for further investigation.

Methods

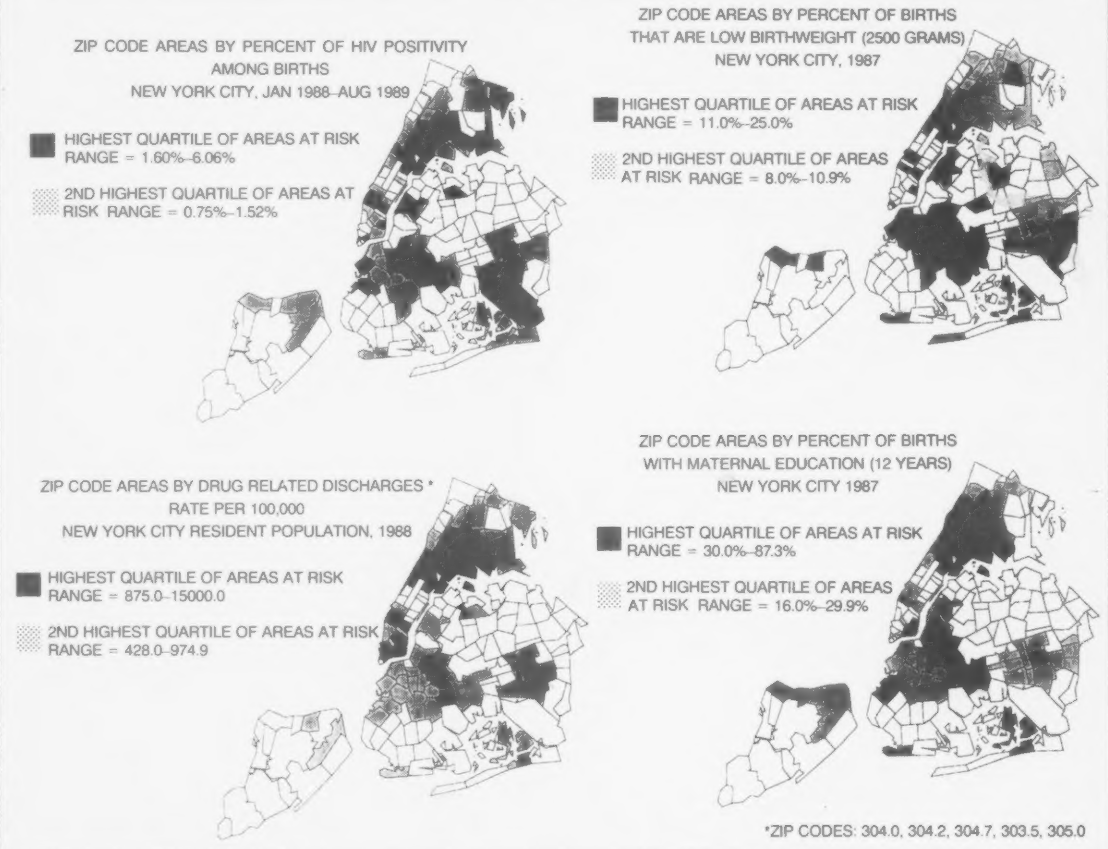
Three sources of data were used: the New York State Newborn HIV Seroprevalence Study, the Birth Certificate Registry, and the New York State hospital discharge file. The data variables analyzed included newborn HIV seroprevalence;[3] hospital discharges with opioid, cocaine, and infant drug withdrawal diagnoses; birthweight; maternal education; and race/ethnicity. All results were converted to rates or percents by zip code, using 1988 postal zip codes.[4]

Newborn HIV seroprevalence and race/ethnicity data included all newborns born after December 31, 1987 and tested for HIV between January 1, 1988 and August 31, 1989. Rates for these variables were calculated as percentages of the total number of newborns tested.

Opioid, cocaine, and infant drug withdrawal diagnoses were obtained from the Statewide Planning and Research Cooperative System for 1988[5] by using the International Classification of Diseases—Ninth Revision codes (304.0, 304.7 and 305.5 for opioid; 304.2 and 305.6 for cocaine; 779.5 and 760.72 for infant drug withdrawal). A combined all-drug variable was established by combining all such adult discharges. Rates per 100,000 population were determined by using zip codes of residence for all persons discharged and zip code population data from the 1985 National Planning Data Corporation estimates.[6] For the infant drug withdrawal rate the denominator was the number of newborn discharges per zip code.

Data on low birthweight (defined as less than 2,500 grams) and low maternal education (defined as fewer than 12 years) were obtained from the registry of 1987 birth certificates.

Geographic comparisons at zip code level were made separately for New York City and for New York State exclusive of New York City. Zip code areas were excluded from the statistical analysis if they had fewer than 100 births, absent variables, or extreme values which were outliers. All variables were mapped by quar-



ZIP CODE AREAS BY PERCENT OF HIV POSITIVITY AMONG BIRTHS NEW YORK CITY, JAN 1988–AUG 1989

ZIP CODE AREAS BY PERCENT OF BIRTHS THAT ARE LOW BIRTHWEIGHT (2500 GRAMS) NEW YORK CITY, 1987

ZIP CODE AREAS BY DRUG RELATED DISCHARGES * RATE PER 100,000 NEW YORK CITY RESIDENT POPULATION, 1988

ZIP CODE AREAS BY PERCENT OF BIRTHS WITH MATERNAL EDUCATION (12 YEARS) NEW YORK CITY 1987

■ HIGHEST QUARTILE OF AREAS AT RISK RANGE = 1.60%–6.06%
▨ 2ND HIGHEST QUARTILE OF AREAS AT RISK RANGE = 0.75%–1.52%
■ HIGHEST QUARTILE OF AREAS AT RISK RANGE = 11.0%–25.0%
▨ 2ND HIGHEST QUARTILE OF AREAS AT RISK RANGE = 8.0%–10.9%
■ HIGHEST QUARTILE OF AREAS AT RISK RANGE = 875.0–15000.0
▨ 2ND HIGHEST QUARTILE OF AREAS AT RISK RANGE = 428.0–974.9
■ HIGHEST QUARTILE OF AREAS AT RISK RANGE = 30.0%–87.3%
▨ 2ND HIGHEST QUARTILE OF AREAS AT RISK RANGE = 16.0%–29.9%

*ZIP CODES: 304.0, 304.2, 304.7, 303.5, 305.0

FIGURE 1—Geographic Comparison of Rates of Newborn HIV Seropositivity, Low Birthweight, Hospital Drug Discharges, and Low Maternal Education by Zip Code, New York City.

tiles for visual ecologic comparison. The geographic distribution of newborn HIV seroprevalence was compared to that of each sociodemographic variable by using Pearson correlation coefficients and multiple regression.

Results

Ecologic Mapping

Visual review of maps of newborn HIV seroprevalence rates by zip code area showed clustering in both New York City (Figure 1) and New York State exclusive of New York City. For New York City all maps of individual drug variables showed a similar pattern so only the most significant variable, all drugs, is presented here. Maps for New York State excluding New York City are not displayed, because release of such data from zip code areas with small numbers could compromise confidentiality.

The first and second quartiles for HIV seroprevalence in New York City had rates of 1.6 to 6.1 percent and 0.7 to 1.5 percent, respectively. The sociodemographic variables were also clustered but appeared to be more closely associated with HIV seroprevalence in New York City than in New York State exclusive of New York City.

Statistical Analysis: New York City

Of the 348 New York City zip code areas listed for 1988, 326 had valid entries listed, but 151 had no population and were excluded, leaving 175. All but five of these areas had a population of 3,798 or more. This restricted file contained 99.5 percent of all births, 99.8 percent of all HIV seropositives, and 99.7 percent of all drug discharges. Further exclusion of nine areas with fewer than 100 births and one area with an extreme drug discharge rate left 165 areas on which the analysis was done. These restrictions entailed minimal loss of data for the variables under consideration (Table IV-1). Over 98 percent of the information for each variable was included in the restricted data set for correlation and regression analyses.

Univariate correlation analysis showed HIV to be correlated in varying degrees with all sociodemographic variables studied in New York City (Table IV-2). Most variables were also highly correlated with one another. If one controlled for low birthweight and the various drug variables in calculating Pearson partial correlation coefficients, low maternal education, percent Black, and percent Hispanic were no longer significantly correlated with HIV. The drug variables were

TABLE IV-1—Number and Percent of Data Points Available for Correlation Analysis by Variable before and after Restrictions, New York City and New York State Exclusive of New York City

Variables	Total Number in Data File	Number Lost To Restrictions*	Number in Analyzed File	Proportional Decrease in Final Analysis
New York City				
Population	7,202,250	37,180	7,165,070	0.5%
Births	196,387	2,732	193,655	1.4%
HIV Positives	2,355	46	2,309	2.0%
All Drug Discharges	47,492	950	46,542	2.0%
Opioid Drug Discharges	28,755	647	28,108	2.3%
Cocaine Drug Discharges	25,701	460	25,241	1.8%
Infant Drug Discharges	4,117	81	4,036	2.0%
New York State Exclusive of New York City				
Population	10,548,363	1,896,471	8,651,892	18.0%
Births	236,648	38,521	198,127	16.3%
HIV Positives	396	54	342	13.6%
All Drug Discharges	12,980	2,179	10,801	16.8%
Opioid Drug Discharges	4,648	877	3,771	18.9%
Cocaine Drug Discharges	9,892	1,575	8,317	15.9%
Infant Drug Discharges	363	40	323	11.0%

*See Text

all highly correlated with one another. A model comprising low birthweight and all drug discharges was shown by multiple regression analysis to be the best predictor of HIV status ($r^2 = .77$). Addition of all remaining variables added little to the model. Low birthweight by itself contributed most to the model ($r^2 = .69$; data available on request).

Statistical Analysis: New York State Exclusive of New York City

Of the 1,918 zip codes listed for New York State exclusive of New York City in 1988, 1,884 had valid entries, but 354 had no population and were excluded. For the remaining 1,530 the median population was 2,441, and 25 percent had populations less than 1,000. This restricted file contained 98.9 percent of all births, 99.2 percent of all HIV seropositives, and 99.2 percent of all drug discharges. Because of small denominators, however, many areas had several extreme or missing values, leading to exclusion of 1,013 areas (1,006 with fewer than 100 births and seven as outliers). These exclusions left only 517 areas for the correlation and regression analyses and resulted in loss of data for a number of variables (Table IV-1). Overall 82 percent of the population, 88 percent of births, 86.4 percent of newborn HIV seropositives, and 83.2 percent of all drug discharges were included in the restricted data set.

Univariate correlation analysis on this restricted data set showed HIV to be less strongly correlated with all variables than in New York City. The variables with the highest correlation were percent Black, percent Hispanic, and the drug variables (Table IV-2). A model comprising percent Black, percent Hispanic, and cocaine drug discharges was shown by multiple regression analysis to be the best predictor of HIV status ($r^2 = .39$). Addition of all remaining variables added little to the model ($r^2 = .43$). The two race variables contributed most to the model ($r^2 = .36$), but all variables were closely associated, and a model without race/ethnicity was almost as good ($r^2 = .42$).

Discussion

The variables used for the analysis were chosen because of their availability, completeness, and suitability as measures of sociodemographic status. Low birthweight has been associated with socioeconomic status.[7] Maternal education, available on the birth certificate, had been demonstrated as a good surrogate for maternal socioeconomic status in previous Departmental studies.[8] Drug discharges, while not a measure of the total drug problem, reflect its distribution by providing a measure of one of its severe outcomes (hospitalization).

Other variables were considered but rejected for one reason or another. For example, sexually transmitted disease rates were not always available by zip code, and Medicaid participation is a biased measure of socioeconomic status because of wide differences in enrollment by locality.

Zip code was chosen as the geographic level of analysis because prevention activities must be targeted to small geographic areas and because a dilutional effect is seen when larger units, such as counties, are used.

Most pediatric AIDS cases are associated with use of intravenous drugs by one or both parents.[1] Of the 590 pediatric cases reported in New York State through October 1989, 346 (59 percent) had mothers who used intravenous drugs, and 108 (18 percent) had mothers who were sexual partners of intravenous drug users.[9] The associations between newborn HIV seroprevalence and hospital intravenous drug discharges in the present study were therefore expected. The high correlation with cocaine discharges, however, is of particular concern. While these mothers may have also used intravenous drugs, the cocaine association could reflect an increased heterosexual transmission of HIV, given the marked increase in syphilis transmission associated with sex in exchange for crack cocaine.[10]

Low birthweight is associated with inadequate health care, poor nutritional status, low socioeconomic status, low level of maternal education, and alcohol and other substance abuse,[7,11-13] including use of cocaine and heroin.[14-16] Thus it was not surprising to find low birthweight closely correlated with newborn seropositivity in New York City. Although drug use, as measured by hospital discharges, added to the association, its contribution was less than that of low birthweight. Hospitalization for drug use measures only the more severe portion of a locality's drug problem. Low birthweight may be a better reflection of the problem for the total community. Theoretically low birthweight could be an outcome of, rather than a surrogate for, HIV seropositivity. However, several studies have shown that there are no significant differences in birthweight between

TABLE IV-2—Correlations between HIV Seropositivity, Low Birthweight, Drug Discharges, Low Maternal Education and Race/Ethnicity, New York City and New York State Exclusive of New York City

	Pearson Correlation Coefficients								
Variables	HIV	LBW	Drug	Opi	Coc	InDr	ME	Blk	His
New York City									
HIV	—	.81	.79	.75	.80	.77	.68	.61	.31
LBW	.81	—	.73	.65	.77	.75	.65	.79	.21
Drug	.79	.73	—	.98	.98	.86	.68	.45	.36
Opi	.75	.65	.98	—	.91	.82	.68	.34	.43
Coc	.80	.77	.98	.91	—	.87	.66	.54	.29
InDr	.77	.75	.86	.82	.87	—	.78	.54	.35
ME	.68	.65	.68	.68	.66	.78	—	.46	.70
Blk	.61	.79	.45	.34	.54	.54	.46	—	−.08
His	.31	.21	.36	.43	.29	.35	.70	−.08	—
NYS Exclusive of NYC									
HIV	—	.33	.49	.41	.50	.35	.33	.50	.43
LBW	.33	—	.44	.33	.46	.35	.52	.55	.23
Drug	.49	.44	—	.84	.98	.37	.44	.67	.47
Opi	.41	.33	.84	—	.72	.23	.31	.44	.46
Coc	.50	.46	.98	.72	—	.41	.50	.70	.45
InDr	.35	.35	.37	.23	.41	—	.30	.34	.21
ME	.33	.52	.44	.31	.50	.30	—	.53	.30
Blk	.50	.55	.67	.44	.70	.34	.53	—	.30
His	.43	.23	.47	.46	.45	.21	.30	.30	—

LBW = Low Birthweight
Drug = All Drug Discharges
Opi = Opioid Discharges
Coc = Cocaine Discharges
InDr = Infant Drug Discharges
ME = Maternal Education < 12 Years
Blk = Percent Black
His = Percent Hispanic

newborns of HIV-positive versus HIV-negative women when the analysis controls for drug use and socioeconomic status.[17-19]

Race/ethnicity was correlated with HIV seroprevalence in New York City, but this correlation appeared to be related primarily to low birthweight and drug use. This finding parallels AIDS case rates, where minorities have much higher case rates but primarily as a function of intravenous drug use.[20]

A different pattern of correlations was seen in New York State exclusive of New York City. The geographic distributions of all sociodemographic variables were still correlated with newborn HIV seroprevalence by zip code, but the correlation coefficients were far less strong. Also, on multiple regression analysis, race/ethnicity replaced low birthweight as the most significant, albeit minimal, predictor of HIV status, while drug use remained a contributor.

Minorities have previously been shown to be disproportionately affected by AIDS, with case rates three to 12 times those of Whites.[21] Blacks and Hispanics contribute 24 percent and 14 percent of the United States AIDS cases, respectively, while comprising only 12 percent and 7 percent of the population.[22] Among AIDS patients, 71 percent of all women are Black or Hispanic, as are 77 percent of all children.[1] These differences have been repeatedly associated with intravenous drug use and secondary heterosexual and perinatal transmission.[20-23]

The reason why race/ethnicity appears to be a better correlate for HIV than low birthweight or drug use in New York State exclusive of New York City is unclear. Some possible explanations are the smaller amount of data available for analysis by zip code, the lower prevalence of drug use resulting in hospitalization (drug discharges per 100,000 population were 17 percent of those in New York City), and the lower rate of newborn HIV seroprevalence (0.16 versus 1.25 percent). Furthermore, the ratios of AIDS case rates for Blacks and Hispanics compared to Whites have been much greater in New York State exclusive of New York City (7.6 and 5.2, respectively, versus 2.1 and 2.0). Similarly, the ratios of HIV newborn seroprevalence rates for Blacks and Hispanics are 15 and 10 times higher, respectively, than for Whites in New York State exclusive of New York City, compared to six and four times higher in New York City.

While the major emphasis for public health HIV/AIDS programs has been on universal education and precautions, efforts targeted at high-risk persons or geographic areas are also important. Our results confirm that prevention activities in New York City should be focused on geographic areas that have low HIV/AIDS rates but populations with low birthweights and high drug use—that is, populations at risk for future HIV/AIDS activity—as well as on geographic areas with already high HIV/AIDS rates. The extent to which sociodemographic factors are important predictors outside New York City remains to be seen.

Summary

The geographic distribution of newborn human immunodeficiency virus seroprevalence at zip code level was compared with the distribution of four sociodemographic variables. For New York City significant univariate correlations were found between HIV and low birthweight, drug use (as measured by hospital discharges), maternal education less than 12 years, and race/ethnicity. Less significant correlations were found for New York State exclusive of New York City. For New York City a model comprising low birthweight and all drug discharges was shown by multiple regression analysis to be most strongly associated with HIV status ($r^2 = .77$). Elsewhere a model comprising race/ethnicity (percent Black, percent Hispanic) and cocaine

drug discharges had the best, but less strong association ($r^2 = .39$). However, because of intercorrelations, race/ethnicity added little when the other variables were included first. Knowledge of the geographic association between HIV seroprevalence and sociodemographic status can be useful in designing and focusing prevention efforts in areas at highest risk for future HIV/AIDS activity. ☐

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V. Projection of AIDS Incidence in Women in New York State

Lawrence Lessner, PhD, MPH

Introduction

One objective of the New York State Newborn HIV Seroprevalence Study was to provide information for AIDS case projections. Earlier projections were developed by linear extrapolation from reported AIDS cases. The limitations of using AIDS case reports include underreporting of AIDS cases, reporting delays, and, most importantly, the long incubation period of HIV infection.

To improve the projection process, we turned to data from two other sources: fertility data, and the results of the Newborn HIV Seroprevalence Study. With these and other data, we estimated the number of HIV-infected women of reproductive age in 1988 and then the number of women infected each year during 1981–88. From this distribution of the incidence of infection and an incubation distribution, we projected the anticipated AIDS case incidence for this population.

The AIDS case projection methods developed here are a form of back calculation.[1,2] The major advantage of this application is that information from another data source, newborn HIV seroprevalence, is employed to obtain an estimate of the number of infected women of childbearing age in 1988.

As newborn serosurveys for antibody to HIV are implemented in more states, this type of analysis will become increasingly important in efforts to anticipate future AIDS cases.

Methods

AIDS case projections were developed in three steps:

1. Newborn seroprevalence results from 1988 provided an estimate of the number of infected women of reproductive age in 1988.

2. An incubation distribution was used to create an expected case series for each number of a family of hypothetical distributions of incident infections. A "best distribution" of infections was selected by comparing the expected and actual case series.

3. A lower bound on future AIDS cases for this population was obtained from the estimates of presently infected women.

Estimating the Number of Infected Women

The number of HIV-seropositive newborns in 1988 was obtained from the Newborn HIV Seroprevalence Study[3] for each of 80 age, racial/ethnic, and geographic strata. The number of HIV-infected women of reproductive age (15–44 years) in each stratum was determined by dividing the number of newborn HIV seropositives[3] by the fertility rate for that stratum. The total estimate is the sum of these 80 stratum-specific estimates.

The fertility rate R of a given female population is defined as B, the number of live births that occur in the population during the given year, divided by P, the average or midyear population for the given female subgroup of age 15 to 44 years: R = B/P. An estimate of P_i, the size of the infected female population, was obtained by inverting the relationship: $R_i = B_i/P_i$. If B_i is the number of seropositive births and R_i is the fertility rate in the infected female population, the population size is $P_i = B_i/R_i$. Assuming the same fertility rate for infected women as for the entire stratum $R_i = R$, then $P_i = B_i/R$.

Since the fertility rate R varies by age, race, and geographic location, the population was stratified by age (15–19, 20–29, 30–39, 40–44 years), race (Hispanic, non-Hispanic Black, non-Hispanic White, and all others), and five geographic regions: New York City (the five bor-

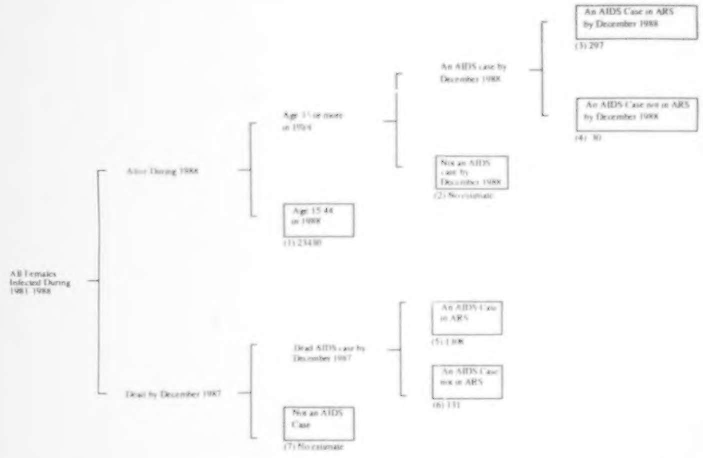
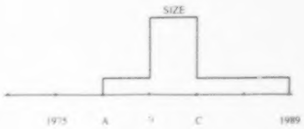

All Females Infected During 1981-1988
Alive During 1988
Dead by December 1987
Age 45 or more in 1988
Age 15-44 in 1988
(1) 23430
An AIDS case by December 1988
Not an AIDS case by December 1988
(2) No estimate
An AIDS Case in ARS by December 1988
(3) 297
An AIDS Case not in ARS by December 1988
(4) 30
Dead AIDS case by December 1987
Not an AIDS Case
(7) No estimate
An AIDS Case in ARS
(5) 1108
An AIDS Case not in ARS
(6) 131

FIGURE 1—Procedure for Estimating the Total Number of Women Infected During 1981–88



A The beginning of the infection epidemic curve, called START, has values 1975, 76, 77, 78, 79, 80

B Called CENTER, is the beginning of the period of high injection activity and has values 1975 to 1980 and is constrained by A≤B

B.C The interval from B to C is called the period of high infection activity, has length C - B = WIDE, has values 1, 2, 3, 4 years

SIZE Is the proportion of cases in the period of high infection activity, has values .5, .6, .7, .8, and .9

FIGURE 2—Generic Candidate For Distribution of Infections

oughs); New York City Vicinity (Nassau, Suffolk, and Westchester Counties), Mid-Hudson Valley (Dutchess, Orange, Putnam, Rockland, Sullivan, and Ulster Counties), Upstate Urban (Albany, Schenectady, Onondaga, Erie, and Monroe Counties), and Upstate Rural (all other counties).

Vital statistics data for 1980 and 1988 were used to obtain estimated stratum-specific fertility rates. The 1980 fertility rates were increased in each stratum to reflect the 1988 fertility experience: ages 15–19 were increased by 17.8 percent, ages 20–29 by 4.1 percent, ages 30–39 by 30.5, and ages 40–44 by 48.9 percent (J. Cooney, Bureau of Production Systems Management, New York State Department of Health, oral communication, January 1990).

The model for estimating the number of women infected during 1981–88 is presented in Figure 1. This tree diagram partitions the ever-infected population into disjoint subgroups for which infection is estimated. Subgroups 4 and 6 use estimates from the AIDS Reporting System (ARS) which is considered 80 to 90 percent complete (Dr. Meade Morgan, Chief of Statistical and Data Management Branch, Division of HIV/AIDS, Centers for Disease Control, oral communication, January 19, 1990).

Estimating the Distribution of Infections

A family of candidate distributions was examined for their utility in describing the incidence of HIV infection in women during 1981–88. The generic candidate (Figure 2), a two-step histogram, was generated by specifying four parameters: START, WIDE, CENTER, and SIZE. START, the beginning of the infection epidemic, had possible values 1975, 1976, . . ., 1980. The interval of high infection activity begins at CENTER and had possible values that range from START to 1988. WIDE is the width of the period of high infection activity and had possible values 1, 2, 3, or 4 years. SIZE is the proportion of infections within the period of intense infection activity and had possible values 0.5, 0.6, 0.7, 0.8 or 0.9. With this model there are 1,200 possible two-step histograms.

A Weibull incubation distribution was selected with a median of 10 years and a fifth percentile of 3.5 years. This selection was based on reports of incubation distributions in a hemophiliac population and in the San Francisco Gay Cohort,[4,5] with estimated median incubation times of 10.0 and 9.8 years, respectively. The estimated fifth percentiles of incubation times were 3.6 and 3.4 years, respectively. An expected case series was created for each of the 1,200 candidates and compared to the annual observed, lag-adjusted AIDS case incidence.

The comparison measure, called ERROR, is the sum of the absolute differences between the observed and expected incidences for each year in each candidate series. The 1,200 ERROR sums were ordered from smallest to largest. Absolute accuracy was measured by ERROR. Relative accuracy was obtained with %ERROR, which is ERROR divided by 2,589, the total incidence observed during 1981–88.

To validate the use of an incubation distribution with a median of 10 years and a fifth percentile of 3.5 years, a family of nine Weibull incubation distributions was created having a 50th percentile of 9, 10, or 11 years and a 5th percentile of 3.0, 3.5 or 4.0 years. Each of these incubation distributions and each of the 1,200 histograms of infection were combined to produce 10,800 candidates, and for each candidate an expected case series was calculated. The top 50 candidates, as ranked by ERROR, were examined to find the smallest ERROR for each incubation distribution (Table V-1). This calculation was at least consistent with the choice of 10 and 3.5 years, although the choices (10, 4.0) and (9, 4.0) appeared equally good. The choices where the 50th percentile is 11 or the 5th percentile is 3.0 appeared considerably less consistent with the observed case series.

Projections of AIDS Case Incidence in Women

The candidate with the minimum value of ERROR was selected as the estimate of the distribution of infections. The expected number of cases was calculated for each year as a lower bound on the anticipated future incidence of AIDS cases in women.

TABLE V-1—Comparison of Incubation Distribution Parameters and ERROR of the Best Candidate

Candidates for the 50th Percentile	Candidates for 5th Percentile		
	3.0	3.5	4.0
9	(*)	70.9	58.9
10	123.7	61.0	61.9
11	121.4	91.2	105.8

(*) at least 154.5

NOTE: Entries are the ERROR for the best candidate for each combination of parameters

TABLE V-2—Summing the Decision Tree (Figure 1)

Box	Population (Origin of the estimate)	The Estimate
1	Women of childbearing age in 1988 (Inverting the fertility rates over 80 strata)	23,430
2	Alive in 1988 and age 45 or more, but not an AIDS case by Dec. 1988 (Unknown)	No estimate
3	An AIDS case in ARS by Dec. 1988 (ARS)	297
4	An AIDS case by Dec. 1988 but not in ARS (10% of 297)	30
5	AIDS case in ARS and dead by 1987 (ARS)	1,308
6	Dead AIDS case by Dec. 1987 and *not* in ARS (10% of 1,308)	131
7	Not AIDS case and died from either other causes or HIV-related disease not AIDS (Unknown)	No estimate
Total		25,196

Results

Estimates of Number of Infected Women (15–44 years)

If we assume equal fertility rates for the general and infected populations in each of 80 age, racial/ethnic, and geographic strata, an estimated total of 23,430 women age 15–44 years were infected with HIV in New York State in 1988. This total was used to obtain an estimate of all women infected during 1981–88 (Figure 1, Table V-2).

Estimates were not possible for certain populations. Box 2 represents all women infected during this period who were alive and age 45 years or more in 1988 but who were not a defined Centers for Disease Control (CDC) case by December 1988. This is the only source of infected women who are potential future AIDS cases for which there is no estimate. Box 7 represents all women infected during this period who died by December 1987 from causes not related to HIV infection. This population will not influence future AIDS incidence.

The estimate of all HIV-infected women of reproductive age in New York State during 1981–88 is 25,196.

Estimates of Distribution of Infections during 1981–88

The 1,200 candidates were ordered by increasing values of ERROR, and the observed and expected cases for six representative candidates were compared (Table V-3). The best candidate had half the ERROR of the next-best candidate.

Three populations of candidates for the infection distribution—the best 10, 20, and 30 candidates—were examined for consensus on the four parameters (Table V-4). START had a clear single mode across all three populations. This is a strong consensus that year 1980 was the beginning of the infection distribution. For CENTER there was a strong consensus across all three populations for 1984 and 1985.

The consensus for the percentiles (Table V-5) shows that the cumulative proportion of new infections nearly doubles in 1984 for each of the three populations. Thus the period of intense infection activity began in 1984. The three-year period 1984–86 accounts for 70 to 80 percent of all infections.

Projections of Future AIDS Cases

The three best candidates for the distribution of infections were used with the incubation distribution to project future AIDS case incidence for 1989 to 1993 (Table V-6). A considerable consensus was found for each year across the three best candidates, and the means and standard deviations for the top 20 candidates substantiated this consensus. The average observed coefficient of variation for these five years was 0.0313. The best candidate indicates that if there were no additional infections after December 1988, the minimum cumulative number of cases expected December 1993 is 8,295.

The linear projections in Table 6 are estimates of future AIDS incidence calculated from actual monthly AIDS surveillance data by linear extrapolation.[6] These linear projections and the best candidate show considerable similarity for the first three years. Beginning in 1992 there is a growing divergence. The linear projections continue to increase each year by a fixed amount, 277 cases. The rate of increase in the candidates is governed by the distribution of infections and, in time, if one assumes no additional infections, there will be a slowing in the rate. Since these projections make no attempt to anticipate future infections, they must be viewed as a conservative estimate.

Discussion

These projections of future AIDS cases among women of reproductive age are based on the results of a newborn HIV seroprevalence study in New York State. An earlier method, linear extrapolation, yielded similar projections for the short term.[6] Projections by extrapolation will always exceed those produced with a method that uses a fixed number of infected because the projections of AIDS cases produced here are based only upon presently infected persons; they must be considered a lower bound of expected cases.

Back calculation is a general class of methods that uses an incubation distribution and the observed case incidence series to estimate the distribution of infections over time.[1,2] The method discussed here, employing two-step histogram, suffers from all the usual disadvantages of back-calculation methods: underreporting of case incidence, limited knowledge of the incubation distribution, and absence of case reports before a certain year (in this case 1981). The key theoretical advantage of back calculation, compared to earlier empiric extrapolation, is its direct quantification of the size of a previously infected population.

The potential limitations of the validity of the work presented here include the assumption of equal fertility between the general and the infected populations in each stratum, the selection of an incubation distribution, the simplicity of the two-

TABLE V-3—Expected Case Series for Selected Best Candidates, Assuming 25,196 Total Infected Woman and a Weibull Incubation Distribution Having 50th Percentile of 10 Years and 5th Percentile of 3.5 Years

Year	Observed	Best	Best 2nd	Best 3rd	Best 10th	Best 20th	Best 30th
1981	9	18	29	32	19	23	21
1982	35	50	58	65	52	62	58
1983	96	99	98	110	104	124	116
1984	155	168	159	168	183	210	196
1985	267	269	262	253	305	319	298
1986	435	436	420	411	476	458	426
1987	658	670	637	644	697	649	595
1988	934	939	901	915	957	889	810
Total	2589	2649	2564	2598	2793	2734	2520
ERROR		61.04	121.8	152.07	204.41	253.97	323.23
%ERROR		2.40	4.7	5.90	7.90	9.80	12.50

step histograms, and the underrepresentation of infections occurring during the survey period, 1988.

A fundamental assumption in relating newborn seropositivity to infections in all women of reproductive age is that HIV-infected women have the same fertility rate as uninfected women in the same age, racial/ethnic, and geographic strata. However, the possible effects of HIV infection challenge this assumption. Intravenous drug use could also reduce fertility and thus low-bias the estimate of infected women of childbearing years.

Few published reports address the potential differences in fertility between the general population and HIV-infected or intravenous-drug-using women. Commonly the comparison is between seropositive and seronegative women who are also or have been intravenous drug users.[7-10] For these groups no differences in fertility have been observed.

If there is a decrease in fertility in HIV-infected women, it may be secondary to associated intravenous drug use. Selwyn[7] compared the pregnancy rate of women in the Bronx in 1985 with a cohort of intravenous drug users enrolled in a methadone maintenance treatment program. The latter group had an estimated decrease in the fertility rate of 4.7 pregnancies per 100 person-years. Thus if there is a decrease in fertility, it is likely to be small (Dr. Peter Selwyn, Department of Epidemiology and Social Medicine, Montefiore Medical Center, New York, NY, oral communication, January 1990).

To investigate the sensitivity of the assumption of equal fertility, we computed the estimate of total infected women for progressively decreasing fertility rates in the infected population. If

$$\sum_{i=1}^{80} \frac{B_i}{R_i}$$

is the original estimate of the total number of infected women of childbearing age in 1988,

$$\sum_{i=1}^{80} \left[\frac{B_i}{(1-a)R_i} \right] = \left[\frac{1}{1-a} \right] \sum_{i=1}^{80} \frac{B_i}{R_i}$$

is the estimate resulting from a 100a percent decrease in fertility. The results of these calculations are given in Table V-7, as well as the ERROR for the best, the best five, and the best 10 candidates for each scenario of decreased fertility. These ERROR values measure the consistency of the assumed decrease in fertility.

As fertility decreases, the estimated total number of infected women increases to 48,626 (corresponding to a 50 percent decrease in fertility). The three ERROR values generally increase with decreases in fertility. This generally increasing ERROR implies a degradation in the consistency of the model and casts doubt on the estimated total number of infected women that results from a severe decrease in fertility. Modest decreases in fertility result in larger but still credible estimates of infecteds (a 15 percent decrease in fertility results in a 16 percent increase in infecteds and a 13 percent increase in best-10 ERROR). Larger decreases in fertility are clearly much less credible than the estimate of no difference in fertility rates between the HIV-infecteds and the general population.

The selection of the incubation distribution was based upon two articles[4,5] that indicated consensus in the choice of a Weibull distribution and its parameters.

TABLE V-4—Consensus for Distributional Parameters in the Best 10, 20, and 30 Candidates

Variables	Best 10	Best 20	Best 30
START			
75	1	1	1
76	0	1	1
77	0	2	2
78	0	2	5
79	3	4	7
80	6	10	14
SIZE			
0.5	1	2	2
0.6	3	3	5
0.7	2	5	10
0.8	3	7	9
0.9	1	3	4
WIDE			
1	2	6	8
2	2	3	5
3	4	7	12
4	2	4	5
CENTER			
84	5	10	12
85	4	8	14
86	1	2	4

TABLE V-5—Consensus for Percentiles of Incident Infections in the Best 10, 20, and 30 Candidates

Year	Best 10	Best 20	Best 30
1980	.06 (.01)	.06 (.01)	.06 (.02)
1981	.10 (.02)	.10 (.03)	.10 (.03)
1982	.15 (.04)	.13 (.04)	.14 (.04)
1983	.19 (.06)	.17 (.06)	.18 (.06)
1984	.34 (.07)	.32 (.09)	.30 (.09)
1985	.61 (.12)	.67 (.16)	.63 (.17)
1986	.87 (.09)	.88 (.10)	.86 (.11)

NOTE: The average cumulative number of new infections from 1975 to the given year for the Best 10, 20 and 30 Candidates is shown. The standard deviation for the data is given in parentheses

Attempts at verification were consistent with those reported in both articles.

Although structurally more complex histograms would yield more accurate results, our simple histograms still offer a wide variety of choices. In any case, parametric modelers performing a back calculation with maximum likelihood methods must necessarily limit the form of their candidate populations for each estimate.[2] The accuracy of the expected case series in the best several candidates (Table V-3), the strength of consensus in parameter estimates (Table V-4), and the strength of consensus in the future num-

TABLE V-6—Projections of Future AIDS Cases from Seroprevalence Data and by Linear Extrapolation

		Seroprevalence Projections				
					Summary of the Top 20 Candidates	
Year	Linear Projections*	Best	2nd Best	3rd Best	Mean	SD
1989	1170	1221	1180	1199	1232	47.5
1990	1446	1487	1447	1469	1502	53.0
1991	1723	1714	1679	1702	1732	54.9
1992	1999	1885	1858	1880	1904	52.7
1993	2276	1988	1970	1989	2006	46.6
Total 1989–93	8614	8295	8134	8239	8376	

*Estimates of future AIDS incidence from an empiric model using actual monthly AIDS surveillance data.

TABLE V-7—Sensitivity and Consistency of the Estimate of the Total Number of Women Infected during 1981 to 1988 to Decreases in the Fertility Rate of Infected Women

% Decrease in Fertility of Infected Women	# Women of Childbearing Years	Total Number Infected during 1981–1988	ERROR		
			Best	Best 5	Best 10
0	23430	25196	61	134	164
10	26033	27799	107	144	190
12	26565	28331	71	140	180
15	27565	29331	94	151	185
17	28118	29884	125	162	195
20	29288	31054	175	209	229
22	29897	31663	176	222	237
30	33472	35238	88	175	242
42	40217	41983	206	246	289
50	46860	48626	254	328	363

bers (Table V-6) confirm the adequacy of the simple histograms.

This method, utilizing HIV seroprevalence data from newborns, has resulted in improved projections of AIDS case incidence in women and may be useful in developing improved projections of pediatric AIDS.

Summary

The results of the Newborn HIV Seroprevalence Study in New York State for 1988 were used to estimate the number of human immunodeficiency virus (HIV)-infected women of childbearing age in 1988. The estimate was accomplished for each of 80 age, racial/ethnic, and geographic strata by dividing the number of seropositive newborns by the 1988 estimated fertility rate. Summing across strata yielded a total of 23,430 infected women age 15 to 44 years in 1988. From this estimate a tree model was used to estimate the number of women infected during 1981–88. Back calculation was used to distribute these infections over time, and the selected incubation distribution was used to estimate future AIDS cases for this population: 8,295 cases by 1993. □

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VI. HIV Infection among Women Entering the New York State Correctional System

Perry F. Smith, MD
Jaromir Mikl, MSPH
Benedict I. Truman, MD, MPH
Lawrence Lessner, PhD, MPH
J. Stan Lehman, MPH
Roy W. Stevens, PhD
Elaine A. Lord, MA
Raymond K. Broaddus, PhD, MA
Dale L. Morse, MD, MS

Introduction

Human immunodeficiency virus (HIV) infection is becoming a leading medical problem among prison inmates in the United States, especially in states where the infection is prevalent. By October 1988 a total of 3,136 AIDS cases had been reported among inmates in a survey of 70 federal, state, and local correctional systems.[1] In New York State, which has more AIDS cases than any other state,[2] 1,046 cases had been reported among inmates of the state system by September 30, 1989.[3]

Because of the continuous turnover of prisoners, it is difficult to determine the precise number of persons with AIDS in New York State prisons at any time. However, a survey by the New York State Department of Correctional Services estimated that in September 1989 there were more than 800 inmates with symptomatic HIV disease and that approximately 550 inmates were on zidovudine therapy (Dr. Robert Greifinger, New York State Department of Correctional Services, personal communication). AIDS and HIV-related illness became the leading causes of death among inmates in 1985 and accounted for 68 percent of inmate deaths in 1987.[4]

In addition to severe HIV disease, there is evidence that many New York inmates are asymptomatically infected with HIV. A 1987 study of men entering the main intake center found that 17.4 percent were HIV-seropositive, the highest rate reported in any blind seroprevalence study of prisoners.[1,5] For intravenous drug users the rate was 44 percent. This study suggested that approximately 2,200 of the 12,600 men who entered this center during 1988 were infected with HIV.

HIV infection is also a problem among female prisoners in New York. Females are 4 percent of all inmates and have accounted for 4 percent of all inmate cases (41 cases). Thus the cumulative incidence rate from 1981 through 1988 has been equal for male and female inmates (27 cases per 1,000 inmates).[4]

Although the prevalence of HIV infection among female prisoners had never been studied, the statewide newborn HIV seroprevalence data suggested a major HIV problem among women in New York. This study was initiated to assess the extent of and risk factors for HIV infection among women entering the state prison system. Secondary objectives were to determine the association of HIV infection with clinical markers for syphilis, hepatitis B, and tuberculosis and to compare the fertility of seropositive and seronegative women.

Background

The New York State correctional system is the third largest in the United States, with 56 facilities housing 44,637 inmates at the end of 1988 (Mr. Gerald Bala, New York State Department of Correctional Services, personal communication). Of the approximately 1,800 female inmates, 69 percent come from New York City and 8 percent from the four counties closest to New York City. Forty-seven percent of female inmates are Black, 34 percent Hispanic, and 19 percent White. Nearly half (46 percent) have been convicted of drug-related offenses. The median duration of incarceration for women is 24 months (range, 12 to 900 months).

During 1988 a total of 1,017 new female inmates entered the New York State system. All female inmates enter at one central processing facility, where a medical evaluation is completed. The intake process includes a complete medical history, physical examination, and routine laboratory screening tests. Through 1988,

TABLE VI-1—Characteristics of Women Entering New York State Prisons, by Study Inclusion Status, September-December 1988

Characteristics	Included in Study (N=480)	Excluded from Study (N=40)	Difference (95% Confidence Interval)
Mean Age (Years)	29.4 (479)*	29.6 (40)*	+ 0.2 (− 2.6, 2.2)
Mean Education (Years)	10.2 (472)	10.4 (40)	+ 0.2 (− 0.8, 0.4)
Race/Ethnicity			
% Black	44.8 (215/480)	52.5 (21/40)	+ 7.7 (− 8.4, 23.8)
% Hispanic	37.5 (180/480)	32.5 (13/40)	− 5.0 (−20.1, 10.1)
% White	17.7 (85/480)	15.0 (6/40)	− 2.7 (−14.3, 8.9)
% With Residence in New York City	74.6 (344/461)	84.6 (33/39)	+10.0 (− 2.0, 22.0)
% Intravenous Drug Users	28.6 (136/475)	47.6 (10/21)	+19.0 (− 2.8, 40.7)

*For mean age and education, the numbers in parentheses indicate the number of women with available information. Fractions in the parentheses for the other variables indicate the number of women with each characteristic over the total number for whom information was available.

inmates were tested for HIV antibodies only when there was a medical indication, such as clinical evidence of immunodeficiency. In 1989 the testing policy was broadened to encourage all inmates, especially those with known risk factors, to undergo voluntary testing, regardless of their clinical status, so that appropriate counseling and medical therapy could be provided. The system does not segregate inmates on the basis of their HIV serologic status.

Methods

All women entering the New York State correctional system from September through December 1988 were included in this study. During this period no woman entered the system more than once, ensuring that there were no duplicates. Demographic information was abstracted from computerized prison records, and selected information that was routinely collected at prison intake regarding medical history, fertility, intravenous drug use, and other HIV risk factors was abstracted from each woman's prison medical record by trained personnel from the State Departments of Health and Correctional Services. After routine screening for syphilis, hepatitis B, and tuberculosis was performed and the results were abstracted to the data collection forms, information that identified specific inmates was removed from the forms, and a unique random number was used to link each form with the blood samples for HIV serologic testing. This process assured that the HIV test results could not be linked to specific inmates.

Syphilis screening was performed with the rapid plasma reagin test and, for confirmation of reactive specimens, the fluorescent treponemal antibody-absorption (FTA-ABS) test. Hepatitis B surface antigen (HBsAG) was determined by radioimmunoassay. These tests were performed by standard techniques at a commercial laboratory. Tuberculosis screening was performed with the intradermal tuberculin purified protein derivative Mantoux skin test (5 tuberculin units). Induration of 10 mm or greater was considered a positive reaction.

HIV testing was performed by the New York State Wadsworth Center for Laboratories and Research. Specimens were first tested in a single standard HIV-1 enzyme-linked immunosorbent assay (ELISA) (DuPont). Specimens with color development less than test cutoff values were recorded as ELISA-negative and were not tested further. ELISA-positive specimens were examined in a standard Western blot confirmatory assay (DuPont), and seropositivity was determined according to criteria recommended by the Association of State and Territorial Public Health Laboratory Directors and the Centers for Disease Control.[6]

The data from the forms and results of HIV serologic testing were merged and analyzed with computer software from SAS Institute, Inc.,[7] BMDP Statistical Software, Inc.,[8] and the Centers for Disease Control.[9] Chi-square, Fisher exact, and *t* tests were applied to univariate analysis. Standard methodology was used to calculate the 95 percent confidence intervals for differences in proportions[10] and differences in means.[11] Multivariate analysis to assess risk factors for HIV seropositivity utilized logistic regression. For this analysis women who reported having had sex with an intravenous drug user, bisexual male, or persons with AIDS or HIV infection or who had exchanged sex for money or drugs were classified as having practiced high-risk sex. Potential confounding of the estimated mean difference of live births and pregnancies between seropositive and seronegative women was investigated with a one-variable and two-variable stratified analysis.[12]

Results

During the four-month study period, 520 women entered the New York State prison system through its central processing facility. Of these, 480 women comprised the study group, while 25 were excluded because of missing demographic and medical information and 15 because of insufficient blood for HIV testing. Compared with the study group, a greater proportion of the 40 excluded women were Black, residents of New York City, and/or intravenous drug users (Table VI-1).

HIV Status and Risk Behaviors of the Study Group

In the study group 90 (18.8 percent) women were seropositive by ELISA; all were confirmed positive by Western blot. Maximum and minimum seroprevalence rates were estimated for all 520 women. Serologic results were available for 25 of the 40 excluded women: four were confirmed positive, one was positive by ELISA but indeterminant by Western blot, and 20 were negative. Thus the maximum estimated number of seropositive women was 110 (21.2 percent) of the 520 entrants: 90 in the study group plus four seropositives, one indeterminant, and 15 with insufficient blood for testing in the excluded group. If the 15 women with insufficient blood for testing and the woman with indeterminant results are assumed to be seronegative, the minimum estimated number of seropositive women was 94 (18.1 percent).

Of the 90 seropositive women in the study group, 53 were Hispanic, 31 Black, and six White. Eighty were between the ages of 20 and 39 years, and 82 reported New York City as their residence at the time of arrest. Sixty-one (67.8 percent) reported a history of intravenous drug use, and an additional six (6.7 percent) reported sexual contact with persons at risk for HIV infection. The remaining 23 reported no recognized risk behaviors for HIV infection, although six had a positive syphilis serology and another reported exchanging sex for money or drugs (suggesting prior high-risk sexual activity). An ad-

TABLE VI-2—HIV Seropositivity of Women Entering New York State Prisons by Demographic Characteristics, Potential Risk Factors, and Clinical Markers for Syphilis, Hepatitis B, and Tuberculosis, September-December, 1988.

Demographic, Risk Factor, and Clinical Marker Variables	Number Tested* (N=480)	Number HIV Positive (%) (N=90) (18.8%)	Odds Ratio** (95% Confidence Interval)
Demographic Characteristics			
Age (years)			
17–19	17	2 (11.8)	Referent
[illegible] 29	259	40 (15.4)	1.4 (0.3, 12.8)
30–39	160	40 (25.0)	2.5 (0.5, 23.4)
40–49	34	6 (17.6)	1.6 (0.2, 18.0)
50+	9	1 (11.1)	0.9 (0.0, 20.8)
Race/Ethnicity			
White, not Hispanic***	85	6 (7.1)	Referent
Black, not Hispanic	215	31 (14.4)	2.2 (0.9, 6.8)
Hispanic	180	53 (29.4)	5.5 (2.2, 16.3)
Education (years)			
≥13	14	1 (7.1)	Referent
10–12	313	52 (16.6)	2.6 (0.4, 112.2)
≤9	145	36 (24.8)	4.3 (0.6, 187.5)
Residence at Time of Arrest			
New York State, Excluding New York City	117	6 (5.1)	Referent
New York City	344	82 (23.8)	5.8 (2.4, 16.7)
Bronx	62	16 (25.8)	6.4 (2.2, 21.1)
Brooklyn	77	24 (31.2)	8.4 (3.1, 26.3)
Manhattan	120	32 (26.7)	6.7 (2.6, 20.4)
Queens	68	10 (14.7)	3.2 (1.0, 11.2)
Richmond	17	0 (0.0)	0 (0.0, 6.1)
Birthplace			
USA—Mainland	377	62 (16.4)	Referent
Caribbean	81	26 (32.1)	2.4 (1.3, 4.2)
South and Central America	15	2 (13.3)	0.8 (0.1, 3.6)
Potential Risk Factors			
Intravenous Drug Use			
No	339	27 (8.0)	Referent
Yes	136	61 (44.9)	9.4 (5.4, 16.4)
Sex with IVDU			
No	378	53 (14.0)	Referent
Yes	94	36 (38.3)	3.8 (2.2, 6.5)
Sex with Bisexual Male			
No	452	83 (18.4)	Referent
Yes	18	5 (27.8)	1.7 (0.5, 5.3)
Sex with Person with AIDS/HIV			
No	468	88 (18.8)	Referent
Yes	4	1 (25.0)	1.4 (0.0, 18.2)
Exchange of Sex for Money or Drugs			
No	426	75 (17.6)	Referent
Yes	47	13 (27.7)	1.8 (0.8, 3.7)
Blood Transfusion†			
No	469	85 (18.1)	Referent
Yes	9	5 (55.6)	5.7 (1.2, 28.9)
Evidence of Tattoo			
No	360	54 (15.0)	Referent
Yes	113	32 (28.3)	2.2 (1.3, 3.8)
Clinical Markers			
Syphilis Serology			
Negative	394	60 (15.2)	Referent
Positive	71	24 (33.8)	2.8 (1.6, 5.2)
Hepatitis B Surface Antigen			
Negative	420	78 (18.6)	Referent
Positive	10	3 (30.0)	1.9 (0.3, 8.5)
Tuberculin Skin Test			
Negative	401	77 (19.2)	Referent
Positive	65	13 (20.0)	1.1 (0.5, 2.1)

*Numbers do not always total to 480 due to missing information.

**Odds ratio for HIV seropositivity among women in each subgroup compared with the referent subgroup.

***Includes one Native American

†Since the dates of blood transfusions were not available, all women who had ever received a blood transfusion were considered to have this risk factor.

ditional 11 reported using street drugs nonintravenously, and two others had received tattoos. The remaining three did not report any of these behaviors nor have reactive syphilis serologies.

Associations of Risk Factors with HIV Seropositivity

HIV seroprevalence varied with demographic characteristics (Table VI-2). Women ages 30 to 39 years had the highest rate (25.0). Compared with the rate for Whites, seroprevalence was twice as high among Blacks and four times higher among Hispanics. Seroprevalence increased with decreasing levels of education. Women who resided in New York City at the time of arrest had higher rates than women from the rest of the state. Nearly one-third of women born in Caribbean countries were seropositive.

HIV seropositivity was associated with several potential risk factors for which information was available (Table VI-2). Of the 136 women who reported a history of intravenous drug use, 45 percent were HIV-seropositive, compared with 8 percent of the 339 women who denied intravenous drug use (odds ratio = 9.4). Seroprevalence was also strongly associated with a history of sex with an intravenous drug user, a history of blood transfusion, and evidence of a tattoo on physical examination.

Overall, of women with available test results, 15.3 percent had positive syphilis serology, 2.3 percent were positive for HBsAG, and 13.9 percent had positive tuberculin tests. Of these three clinical markers, a positive syphilis serology was most strongly associated with HIV seropositivity (Table VI-2).

For 472 women, information was available regarding their use or nonuse of specific street drugs (Table VI-3). Among the women denying any drug use, 7.5 percent were seropositive, compared with 8.2 percent of the women who used drugs nonintravenously and 45.9 percent of the women who admitted to intravenous drug use. Among intravenous drug users, those who used cocaine but not heroin had the lowest HIV seroprevalence (25.0 percent), users of heroin and not cocaine had an intermediate seroprevalence (43.6 percent), and users of both cocaine and heroin had the highest seroprevalence (51.3 percent). The corresponding percentages among drug users who denied intravenous use were lower but showed the same increasing trend from "cocaine, no heroin" to "cocaine and heroin."

TABLE VI-3—HIV Seropositivity of Women Entering New York State Prisons by History of Intravenous Drug Use and Type of Drug Used, September-December 1988.*

IV Drug Use Status and Type of Drug Used	Number Tested	Number HIV Positive (%)	Odds Ratio (95% Confidence Interval)	
Women Denying any Drug Use	107	8 (7.5)	Referent	
Women Using Drugs				
Non-Intravenously Only	232	19 (8.2)	1.1	(0.4, 2.9)
Cocaine, no Heroin	104	3 (2.9)	0.4	(0.1, 1.6)
Heroin, no Cocaine	30	5 (16.7)	2.5	(0.6, 9.4)
Cocaine and Heroin	19	4 (21.1)	3.3	(0.6, 14.1)
Drugs other than Cocaine and Heroin	79	7 (8.9)	1.2	(0.4, 3.9)
Women Using Drugs				
Intravenously	133	61 (45.9)	10.5	(4.5, 25.4)
Cocaine, no Heroin	16	4 (25.0)	4.1	(0.8, 18.2)
Heroin, no Cocaine	39	17 (43.6)	9.6	(3.4, 28.0)
Cocaine and Heroin	78	40 (51.3)	13.0	(5.3, 33.4)

*Eight women (three who admitted to intravenous drug use) were excluded because of incomplete or inconsistent information.

TABLE VI-4—Variables Associated with HIV Seropositivity by Logistic Regression Analysis* for Women Entering New York State Prisons, September-December 1988

Independent Variables	Adjusted Odds Ratio** (95% Confidence Interval)	
Intravenous Drug Use	9.4	(4.8, 18.3)
Race/Ethnicity		
Hispanic	5.8	(1.8, 18.0)
Black	3.3	(1.0, 10.6)
Positive Syphilis Serology	3.2	(1.6, 6.7)
New York City Residence	3.2	(1.1, 9.2)
Blood Transfusion***	2.3	(0.4, 14.5)
Evidence of Tattoo	1.6	(0.8, 3.1)
High Risk Sex†	1.3	(0.7, 2.6)

*The regression model included the seven variables shown, as well as age, as a continuous variable.

**The referent group for Hispanic and Black women was White women, and for women with New York City residence was women with residence in New York State excluding New York City.

***Since the dates of blood transfusions were not available, all women who had ever received a blood transfusion were considered to have this risk factor.

†Women who reported having had sex with an intravenous drug user, bisexual male, person with AIDS or HIV infection, or who exchanged sex for money or drugs were classified as having practiced high risk sex.

TABLE VI-5—Fertility of Women Entering New York State Prisons by HIV Serologic Status, September-December 1988

Measure of Fertility	HIV Positive Women	HIV Negative Women
Mean Number of Pregnancies/Woman* (range)	2.7 (0–8)	2.7 (0–13)
Mean Number of Live Births/Woman (range)	1.8 (0–6)	1.7 (0–8)
Fraction of Women Nulliparous (%)	20/90 (22.2)	113/390 (30.3)

*Two seropositive and 15 seronegative women with incomplete pregnancy histories were excluded from this analysis.

Multivariate analysis identified five factors associated with HIV seropositivity (Table VI-4): women who were intravenous drug users, of Hispanic or Black race/ethnicity, with a positive syphilis serology, or residents of New York City were more likely to be seropositive than women without these factors. There were no interaction effects among any of the variables.

Fertility

Of 21 women who were known to be pregnant, two (9.5 percent) were HIV positive. The mean number of pregnancies, the mean number of live births, and the percentage of women who were nulliparous were similar for HIV-positive and HIV-negative women (Table VI-5). Stratified analysis to control for differences in age, racial composition, and intravenous drug use between seropositive and seronegative women did not change these findings (data not shown).

Discussion

This study, as well as an earlier study of male prison entrants,[5] was crucial in defining the magnitude of the HIV epidemic among New York State prisoners, describing its epidemiology, and promoting increased public health attention to address the epidemic in this high-risk population.

HIV seroprevalence was higher among male (17.4 percent) and female (18.8 percent) entrants to New York State prisons than has been previously reported in any mass screening program or blind epidemiologic study among prisoners in the United States.[1] The seroprevalence among all 520 female entrants during the study period may have actually been closer to our maximum estimate of 21 percent, because many of the excluded women with insufficient blood samples for HIV testing may have had scarred veins from intravenous drug use and been at increased risk for HIV infection. Most seroprevalence rates reported among prisoners in other jurisdictions have been less than 3 percent, with the notable exception of high rates for males (7.0 percent) and females (15.4 percent) entering Maryland State prisons in 1985.[1] The Maryland study found that seroprevalence among male entrants was constant from 1985 to 1987.[13] Follow-up studies of New York State inmates are planned to determine trends in seroprevalence with time.

The blinded design of our study precluded collection of extensive risk information, and the cross-sectional design prevented assessment of temporal associations between HIV infection and risk factors or other infectious diseases. However, the high rate of HIV seropositivity among female inmates is consistent with their relatively high frequency of known risk behaviors, especially intravenous drug use, and the fact that most of them come from New York City, where HIV infection is prevalent. In addition to intravenous drug use, we found that a positive syphilis serology was strongly associated with HIV positivity. Of the 23 HIV-positive women who denied any known risk behavior, six had a positive syphilis test. The association between genital ulcer disease and HIV infection has been previously reported.[14–16] Genital ulcer disease may be a marker for prior high-risk sexual activity and may increase the risk of HIV transmission by disrupting the vaginal or cervical mucosa.

Although HIV infection has not been shown to be associated with tattooing,[17,18] HIV transmission through tattoo needles is biologically plausible, and the possible spread of HIV by this route to prisoners has been reported.[19] In our study, women with tattoos were twice as likely to be HIV-seropositive as women without tattoos (Table VI-2). However, when other possible exposures were controlled in the multivariate analysis, tattoos were not associated with HIV positivity, suggesting that tattooing was not a major factor for transmission in this population.

Our study showed an association between high HIV seroprevalence and Black or Hispanic ethnicity. This association has been described previously among intravenous drug users and attributed to differences in the social setting of drug use and specific drug use behaviors for persons of different ethnic backgrounds.[20–22] Our results may have been affected by these factors, as well as by other unmeasured risk behaviors, such as sexual contact with a person not recognized as being at high risk for HIV infection.

Although other studies have found that intravenous cocaine use is more strongly associated with HIV infection than heroin use,[21,23] our results suggested the opposite (Table VI-3). However, the small number of women in our study limited its power to detect differences by drug use. In any case, caution is needed in interpreting these results, since we had no information on the frequency, quantity, or pattern of intravenous drug use, or on the frequency of high-risk sexual activity.

HIV-positive women in our study reported the same number of pregnancies and live births as the seronegative women, and similar percentages of each group were nulliparous (Table VI-5), even after controlling for differences in age, race, and intravenous drug use between the two groups. These findings suggest that HIV infection did not affect fertility, but because we had no information on the timing of HIV seroconversion relative to the reported pregnancies, the exact relationship between HIV infection and fertility was impossible to assess. Our results are consistent with those of a study among New York City intravenous drug users which found that asymptomatic HIV infection was not associated with a decreased pregnancy rate or an increased risk of adverse pregnancy outcomes.[24] Another study of 136 pregnancies among British women also found no evidence that HIV infection affected the outcome of pregnancy, although there were more spontaneous abortions among the seropositive women.[25]

Determining the impact of HIV infection on fertility would provide important information for interpreting the blind HIV newborn seroprevalence studies that are being conducted widely. These studies provide nearly complete testing of all newborns and therefore are useful measures of HIV-infection among women giving birth and, by extrapolation, among all women.[26] However, if HIV infection substantially reduces fertility, infected women would be less likely to deliver babies, and newborn seroprevalence studies would underestimate infection rates among women. Current evidence suggests that asymptomatic HIV infection may not reduce fertility, but additional studies are needed.

After our study, funding was increased to enable four teams of counselors to provide prisoner education regarding HIV infection and to encourage all prisoners to consider HIV antibody testing. Our study and the earlier study of male inmates[5] also helped to quantify the need for increased medical services for HIV-infected prisoners. Applying the seroprevalence rates from these two studies to the number of entrants at the two prison centers, we estimate that in 1988 approximately 2,200 men and 200 women entered New York State prisons with HIV infection. In addition to routine medical care, these inmates need medical evaluation for newly approved therapies for asymptomatic HIV infection and adequate follow-up of other associated medical conditions, especially tuberculosis, which has shown an alarming increase in association with HIV infection.[27] The problem of perinatal transmission in prison was also apparent from our study, since two of the 21 pregnant women were HIV positive. To address these needs, provision for increased medical care is being planned in New York State prisons; additional seroprevalence studies will be conducted to monitor the HIV epidemic among prisoners in the future.

Summary

Human immunodeficiency virus infection is the leading medical problem among prison inmates in several states. In 1988 a blinded seroprevalence study was conducted on 480 New York female prison entrants to determine the prevalence of and risk factors for HIV infection in this population. Ninety (18.8 percent) women were HIV-seropositive. Seroprevalence was highest among women ages 30–39 (25.0 percent) and varied by ethnicity (Hispanics, 29.4 percent; Blacks, 14.4 percent; Whites, 7.1 percent) and residence (New York City, 23.8 percent; Upstate, 5.1 percent). Nearly half (44.9 percent) of the 136 acknowledged intravenous drug users and one-third (33.8 percent) of the 71 women with a positive syphilis serology were HIV-seropositive. There was no difference in fertility histories between seropositive and seronegative women, and two of 21 pregnant women were seropositive. This study led to increased clinical and prevention services for this high-risk population. ☐

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VII. HIV Seroprevalence in Clients of Sentinel Family Planning Clinics

Rachel L. Stricof, MPH
Thomas C. Nattell, MA, MSW
Lloyd F. Novick, MD, MPH

Introduction

Family planning clinics serve predominantly young, sexually active clients, a substantial proportion of whom may be infected—or at risk of infection—with human immunodeficiency virus (HIV). The purposes of this study were to document the prevalence of HIV antibody in women attending sentinel family planning clinics and to monitor the trends in infection over time.

In New York State, there are approximately 200 discrete family planning clinics, located in all counties except Putnam (where negotiations for a clinic are in progress). All but one of these clinics are publicly subsidized. Other groups and/or individuals provide family planning services but d[illegible] have a clinic dedicated solely for t[illegible]pose.

Fam[illegible] planning clinics provide medically approved methods of contraception, medical evaluations, screening and treatment for vaginal infections and sexually transmitted diseases, level-I screening for infertility, and voluntary HIV counseling and testing services.

Most clients in New York State are women. Approximately 75 percent are under age 25. Overall 21 percent are Black and 13 percent Hispanic. [In New York State approximately 14 percent of the population is Black and 12 percent of Hispanic origin (Current Population Survey: March 1988)].

One constraint on this study was that relatively few of the clinics routinely obtain blood specimens. Random sampling of clients and/or clinics was therefore not feasible. Instead the qualifying clinics were considered sentinels, and the results were adjusted for extrapolation to the total population of clients.

Methods

The study was conducted over a period of 26 months, from February 1, 1988 through March 31, 1990.

The study sites were limited to publicly subsidized family planning clinics that routinely drew whole blood specimens for syphilis serology from clients during the initial medical evaluation. Clinics throughout the state were screened for this practice. In early 1988, 38 clinics met this criterion, and 37 agreed to participate. Of the participating facilities, 13 were located in New York City boroughs and the rest in other counties: Westchester, seven; Nassau, four; Onondaga, three; Monroe, five; Orleans, one; Genesee, one; and Erie, three. One clinic in Monroe County was closed shortly after the study was initiated. Five additional clinics (Rockland, two; Orange, two; Nassau, one) entered the study after changes in their routine blood-collection practices.

Study subjects were limited to clients undergoing an initial medical evaluation for contraceptive services. Clients seen for annual or follow-up visits were excluded to avoid repeat testing of the same individuals over time. Individuals seen solely for pregnancy testing were also excluded, because blood specimens were not routinely obtained on this population. Clients seen only for HIV counseling and testing were excluded to minimize self-selection bias.

If a client met these criteria, a seroprevalence study test requisition (SSTR) form was completed by the clinic staff and submitted along with the routine laboratory request form and blood specimen to a contract laboratory. The completed form contained the client's age, race or ethnicity, and county and zip code of residence, as well as the clinic site and the month and year of specimen collection. Identifying information (e.g., name, street address, or medical record number) was not collected; thus the information could not be traced back to the individual client.

If serum was left over after all routine laboratory tests were performed, the contract laboratory would pipette at least 0.5 ml into a tube labeled with the blind seroprevalence number and send the serum with the SSTR form to the New York State Department of Health's Wadsworth Center for Laboratories and Research. No additional blood was drawn for this study.

When the blood drawn was insufficient for blind HIV testing, the SSTR alone was sent to the Department of Health. These data were used to estimate any bias from the loss of specimens. The data were not used for any other analysis.

The Wadsworth Center performed an ELISA (enzyme-linked immunosorbent assay) (DuPont) for HIV antibody. ELISA reactives were confirmed with a Western blot (in-house assay until May 1988, then DuPont). All Western blots were interpreted by the latest Association of State and Territorial Public Health Laboratory Directors criterion[1] for reactive specimens, i.e. virus-specific bands in at least two of the following major band locations: gp160/120, gp41, or p24. Specimens with other major virus-specific bands were interpreted as equivocal. All other specimens were considered nonreactive.

Seroprevalence rates were calculated as: (number of Western blot reactives)/(total number of specimens minus equivocals). Because the specimens could not be traced to individual clients, we could not request second samples for reevaluation of equivocals.

Age-specific seroprevalence rates were calculated in five-year intervals for ages 15 to 39. Individuals under age 15 were grouped, as were those over age 39. Racial/ethnic-specific seroprevalence rates were calculated for non-Hispanic Whites, non-Hispanic Blacks, Hispanics, Others, and Unknowns. Others and Unknowns had to be combined due to small cell sizes for some combinations of age and race/ethnicity groups.

HIV seroprevalence rates by quarter-year of clinic visit were calculated to describe changes in seroprevalence over time. For this analysis three of the 37 initial clinics were eliminated: one that closed early in the study and two that closed later. Trends were calculated only for the 34 clinics that participated throughout the entire study period.

To estimate the completeness of the data set, we used the number of initial medical examinations in 1988 reported by each participating clinic to the New York State Department of Health's Bureau of Reproductive Health. This number was compared to the number of specimens submitted during the first year of the study (February 1988–January 1989). Clinic records were available from all but one site in upstate New York.

Bureau of Reproductive Health statistics were also used to adjust for differences between the sentinel clinic clients and the overall family planning clinic population. The age, racial/ethnic, and residence distributions of the two groups were compared and the HIV seroprevalence rates were adjusted directly for differences, using the overall population as the standard.

The results of the sentinel clinic study were compared to the results of the Newborn HIV Seroprevalence Study. Sentinel clinic rates were adjusted for the differences in age, racial/ethnic, and residence distributions. The adjusted rates were calculated for New York City, New York State exclusive of New York City, and statewide, using the distribution of childbearing women as the standard population.

Between February and July 1988 a separate, unlinked HIV risk assessment questionnaire was administered to clients of 30 family planning clinics by clinic staff. Participation in the risk study was voluntary and not limited to the subset of women seen for an initial medical evaluation. The results were submitted to the New York State Department of Health without patient identifiers. Participation was voluntary on the part of the facility and the clients.

The questionnaire included items on the client's age; race/ethnicity; county of residence; history of intravenous drug use; history of living in a high-risk country (as defined by the Centers for Disease Control); history of being a sex partner of a bisexual male, drug user, hemophiliac, or person with AIDS or HIV infection; and maximum number of sex partners in any six-month period since 1978 (0, 1, 2, 3, or 4+).

Results

Sentinel Clinic Results

Between February 1, 1988 and March 31, 1990, a total of 28,745 specimens were obtained from women having initial medical examinations in the sentinel family planning clinics. Of these specimens, 896 (3 percent) had an insufficient quantity of serum left over for analysis. An additional 1 percent were not tested due to improper shipping, leakage of tubes, or blank forms or because the specimen was obtained from a male client. Twelve (0.04 percent) had equivocal HIV test results and were eliminated from the analysis. Of the 27,549 remaining specimens 144 (0.52 percent) were HIV antibody-positive.

Fourteen of the 41 clinic sites (one in New York City, 13 in New York State exclusive of New York City), which contributed 11.8 percent of the study population, had no seropositives identified during the study period. Eight clinics (two in New York City) had rates of seropositivity between 0.01 and 0.50 percent; 11 clinics (six in New York City) had rates between 0.51 and 1.00 percent; and eight clinics (four in New York City) had rates in excess of 1 percent. The highest seroprevalence rate was 3.91 percent, detected in a mobile clinic in New York City.

The highest seroprevalence rate by residence of the client was detected among residents of New York City, with 0.82 percent identified as positive. The rates by boroughs (and the numbers tested/received) were: Manhattan, 0.73 percent (15/2,050); Bronx, 0.98 percent (20/2,031); Kings, 0.81 percent (50/6,187); Queens, 0.84 percent (20/2,387); and Staten Island, 0.67 percent (6/891).

Residents living elsewhere in the state had a 0.24 percent rate of seropositivity. Clients from counties nearest New York City (Westchester, Nassau, and Rockland) had five times the seroprevalence of residents from other counties (Erie, Monroe, Orleans, Genesee, Onondaga, and Orange): 0.54 versus 0.11 percent, respectively.

No seropositives were detected among the 805 individuals tested under the age of 15, and three seropositives (0.42 percent) were detected among the 722 individuals age 40 or more (Table VII-1). Between these groups HIV seroprevalence increased with age. Women 35–39 years old had approximately nine times the HIV seroprevalence rate of 15–19 year olds (1.56 versus 0.18 percent). The trend of increasing HIV seroprevalence with age was seen for all racial/ethnic groups.

Minority women (non-Hispanic Black or Hispanic) had an HIV seroprevalence of 0.76 percent (Table VII-1). This was about six times the rate detected among non-Hispanic Whites (0.13 percent).

HIV seroprevalence rates by quarter-year of clinic visit are shown in Table VII-2. The highest rates (0.63 and 0.71

TABLE VII-1—HIV Seroprevalence by Age and Racial/Ethnic Status of Sentinel Family Planning Clinic Clients, February 1988–March 1990

Age (years)	Percent (Number positive/number tested) Non-Hispanic White	Non-Hispanic Black	Hispanic	Other/ Unknown	Total
A. New York City					
<15	0.00 (0/12)	0.00 (0/256)	0.00 (0/123)	0.00 (0/4)	0.00 (0/395)
15–19	0.00 (0/492)	0.25 (9/3,638)	0.30 (6/1,997)	0.00 (0/136)	0.24 (15/6,263)
20–24	0.89 (3/337)	1.07 (19/1,769)	0.88 (11/1,253)	0.00 (0/87)	0.96 (33/3,446)
25–29	1.26 (2/159)	1.92 (14/730)	2.09 (13/622)	1.56 (1/64)	1.90 (30/1,575)
30–34	1.28 (1/78)	2.17 (8/369)	2.26 (9/399)	3.23 (1/31)	2.17 (19/877)
35–39	3.45 (1/29)	1.83 (3/164)	2.05 (5/244)	11.11 (1/9)	2.24 (10/446)
>39	4.76 (1/21)	1.12 (1/89)	0.00 (0/317)	0.00 (0/8)	0.46 (2/435)
Unknown	0.00 (0/3)	0.00 (0/60)	7.41 (2/27)	0.00 (0/19)	1.83 (2/109)
Total	0.71 (8/1,131)	0.76 (54/7,075)	0.92 (46/4,982)	0.84 (3/358)	0.82 (111/13,546)
B. New York State Exclusive of New York City					
<15	0.00 (0/191)	0.00 (0/177)	0.00 (0/27)	0.00 (0/15)	0.00 (0/410)
15–19	0.03 (1/3,816)	0.42 (4/953)	0.27 (1/367)	0.00 (0/204)	0.11 (6/5,340)
20–24	0.04 (1/2,799)	0.46 (4/872)	0.50 (3/606)	0.00 (0/142)	0.18 (8/4,419)
25–29	0.10 (1/1,013)	1.31 (6/459)	0.49 (2/406)	0.00 (0/72)	0.46 (9/1,950)
30–34	0.41 (2/490)	1.93 (4/207)	0.00 (0/237)	0.00 (0/40)	0.62 (6/974)
35–39	0.00 (0/173)	2.47 (2/81)	0.88 (1/114)	0.00 (0/17)	0.78 (3/385)
>39	0.00 (0/131)	1.61 (1/62)	0.00 (0/86)	0.00 (0/8)	0.35 (1/287)
Unknown	0.00 (0/93)	0.00 (0/25)	0.00 (0/37)	0.00 (0/71)	0.00 (0/226)
Total	0.06 (5/8,706)	0.74 (21/2,836)	0.37 (7/1,880)	0.00 (0/569)	0.24 (33/13,991)
*C. Statewide**					
<15	0.00 (0/203)	0.00 (0/433)	0.00 (0/150)	0.00 (0/19)	0.00 (0/805)
15–19	0.02 (1/4,314)	0.28 (13/4,592)	0.30 (7/2,364)	0.00 (0/340)	0.18 (21/11,610)
20–24	0.13 (4/3,138)	0.87 (23/2,641)	0.75 (14/1,859)	0.00 (0/229)	0.52 (41/7,867)
25–29	0.26 (3/1,172)	1.68 (20/1,190)	1.46 (15/1,028)	0.74 (1/136)	1.11 (39/3,526)
30–34	0.53 (3/568)	2.08 (12/577)	1.41 (9/637)	1.41 (1/71)	1.35 (25/1,853)
35–39	0.50 (1/202)	2.04 (5/245)	1.68 (6/358)	3.85 (1/26)	1.56 (13/831)
>39	0.66 (1/152)	1.32 (2/151)	0.00 (0/403)	0.00 (0/16)	0.42 (3/722)
Unknown	0.00 (0/96)	0.00 (0/85)	3.12 (2/64)	0.00 (0/90)	0.60 (2/335)
Total	0.13 (13/9,845)	0.76 (75/9,914)	0.77 (53/6,863)	0.32 (3/927)	0.52 (144/27,549)

*Includes 12 clients for whom no county of residence was identified.

TABLE VII-2—HIV Seroprevalence by Quarter-Year of Clinic Visit for Sentinel Family Planning Clients, February 1988–March 1990

Quarter	Number Positive	Number Tested	Percent Positive
1st 1988	8	2,330	0.34
2nd 1988	19	3,033	0.63
3rd 1988	13	3,162	0.41
4th 1988	18	3,283	0.55
1st 1989	16	3,350	0.48
2nd 1989	13	3,010	0.43
3rd 1989	13	2,710	0.48
4th 1989	13	2,394	0.54
1st 1990	18	2,521	0.71

*Data limited to 34 clinics that participated during the entire study period.

percent) occurred in the second quarter of 1988 and first quarter of 1990. The lowest rate (0.34 percent) was detected in the incomplete first quarter of 1988. No trend over time is evident.

Extrapolation to All Family Planning Clinics

According to Bureau of Reproductive Health statistics, a statewide total of 68,644 initial medical examinations at family planning clinics were done in 1988. Of these examinations, 15,112 (22 percent) involved clients at the sentinel clinics. During the first year of the study (beginning in February 1988) 12,575 client specimens were included in the analysis. This is an estimated sampling rate of 83 percent in the sentinel clinics for 1988: 86 percent in New York City, 80 percent in New York State exclusive of New York City.

Sentinel clinic clients were more likely to be non-Hispanic Black (36.5 percent) or Hispanic (25.3 percent) than were clients in general (21.2 percent and 13.5 percent, respectively). The age distributions were similar: women under age 25 were 72.3 percent of clients at sentinel clinics and 73.4 percent at all clinics statewide.

After adjustments for differences in the racial/ethnic and residence distributions, the extrapolated HIV seroprevalence rates for all family planning clinic clients were: New York City, 0.82 percent; New York State exclusive of New York City, 0.15 percent; and statewide, 0.31 percent.

TABLE VII-3—Risk Factors for HIV Infection of Clients at Sentinel Family Planning Clinics, February 1988–July 1988

Risk Factor	Number Admitting	Number Responding	Percent Admitting
New York City			
IV drug use	13	4,692	0.28
SP* IVDU	47	4,691	1.00
SP* bisexual	35	4,703	0.74
SP* hemophiliac	4	4,680	0.09
High-risk country	28	838	3.34
SP* HIV+	24	4,642	0.52
Having ≥4 SP**	184	4,669	3.94
Any risks above	286	4,703	6.08
New York State Exclusive of New York City			
IV drug use	28	2,174	1.29
SP* IVDU	81	2,179	3.72
SP* bisexual	34	2,189	1.55
SP* hemophiliac	1	2,179	0.05
High-risk country	32	2,176	1.47
SP* HIV+	7	2,173	0.32
Having ≥4 SP**	194	2,141	9.06
Any risks above	297	2,189	13.57
Total			
IV drug use	41	6,866	0.60
SP* IVDU	128	6,870	1.86
SP* bisexual	69	6,892	1.00
SP* hemophiliac	5	6,859	0.07
High risk country	60	3,014	1.99
SP* HIV+	31	6,815	0.45
Having ≥4 SP**	378	6,810	5.55
Any risks above	583	6,892	8.46

*SP—Sexual partner(s).
**SP—Sexual partners in a six-month period since 1978.

Comparison to Newborn Seroprevalence Data

Because there were substantial differences between the demographic profiles of family planning clinic clients and those of childbearing women, adjusted seroprevalence rates were calculated using the age, racial/ethnic, and residence distributions of childbearing women as the standard.

The adjusted seroprevalence rates were: New York City, 1.53 percent; New York State exclusive of New York City, 0.25 percent; and statewide, 0.92 percent. These figures were 24 to 47 percent higher than the rates for childbearing women in the Newborn HIV Seroprevalence Study: New York City, 1.24 percent; New York State exclusive of New York City, 0.17 percent; and statewide, 0.66 percent.[2]

Risk Factors

Thirty sentinel clinics participated in the voluntary risk assessment study. Risk assessment forms were completed on 6,892 clients, who were generally of the same age (71 percent under age 25) as the whole study population. The risk-assessment group was more likely to reside in New York City and to be non-Hispanic Black (50 percent) or Hispanic (27 percent) than the overall sentinel population. (Of the risk-assessment group 21 percent were non-Hispanic White, and 3 percent were of Other or Unknown race/ethnicity.)

The number and percent of clients in the sentinel clinics admitting to the selected risk factors for HIV infection are presented in Table VII-3. The most frequently reported risk factors were having sex with four or more sex partners in a six-month period, having lived in a high-risk country, having had sex with an intravenous drug user, and having had sex with a bisexual male. Of all the participating clients, 583 (8.46 percent) admitted to at least one of the risk factors. Fifty-nine (11.64 percent) of the pregnant women admitted to at least one risk factor.

The women who volunteered for this study were not limited to those in the seroprevalence study. Not all sentinel clinics participated, and the extent of participation varied considerably among the participating clinics. For these reasons the results should be considered background information and not interpreted to explain seroprevalence results obtained for the full study population.

Discussion

Family planning clinics serve predominantly young, sexually active clients. In this study of 41 sentinel clinics, approximately one in 200 clients was infected with HIV.

The clinics in this study were not representative of all of the approximately 200 clinics in the state. Sentinel clinics were limited to those that routinely drew blood for syphilis serology on all initial clients. These clinics tended to be in areas with higher syphilis incidence rates and thus potentially higher HIV infection rates. The crude and adjusted HIV seroprevalence rates obtained in this study may overestimate the actual rate for the total family planning clinic population. The true rates are likely to be closer to those found in the Newborn HIV Seroprevalence Study.

A slightly low bias in statewide rates may have been introduced by the 896 specimens that were insufficient for testing. These specimens represented clients who resided primarily in New York City (79.7 percent). The 896 clients resembled the entire study group in their age distribution (71.5 percent under age 25) and resembled other New York City clients in their racial/ethnic distribution (non-Hispanic White, 21.0 percent; non-Hispanic Black, 42.5 percent; Hispanic, 34.4 percent; Other, 2.1 percent).

After reviewing the results of this and other seroprevalence studies involving New York State residents, the New York State Department of Health required HIV education, counseling, and testing programs as a condition of funding in all publicly subsidized clinics for family planning and prenatal care.

Since virtually all women seen in these clinics are sexually active, they are potentially at risk for HIV infection. Knowledge of an individual's HIV infection status is important to prevent transmission, particularly vertical transmission from mother to infant, and to prevent or defer certain complications of HIV infection with early therapeutic intervention.

In addition to providing HIV counseling and testing, family planning and prenatal care clinics must be prepared to provide or offer referral for care for their HIV-positive clients.

Summary

In February 1988 the New York State Department of Health initiated a study to determine the prevalence of HIV antibody in women attending selected, publicly subsidized family planning clinics. During a 26-month study period, 27,549 blood specimens were obtained from women having an initial medical examination in 41 clinic sites throughout the state. Of these clients 144 (0.52 percent) were seropositive.

The HIV seroprevalence rate increased with age to a high of 1.56 percent for 831 women ages 35 to 39. The seroprevalence rate for non-Hispanic Black or Hispanic clients (0.76 percent) was about six times the rate for non-Hispanic Whites (0.13 percent). No overall increasing or decreasing trend in prevalence of HIV infection was detected during the study period. ☐

Acknowledgments

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VIII. Hospitalization of HIV-Seropositive Newborns with AIDS-Related Disease within the First Year of Life

Donna M. Glebatis, MS
Lloyd F. Novick, MD, MPH
Andrew Stacy, BA

Introduction

HIV seropositivity in a newborn is a reflection of the mother's HIV infection status, not necessarily that of the infant.[1,2] Only a minority of seropositive infants are actually infected, remaining HIV-positive after 18 months. Studies of perinatally acquired infection have shown wide variation in the rates of maternal transmission.[3-10] Factors such as low numbers of study subjects, loss to follow-up, neonatal deaths from indeterminate causes, absence of detectable antibodies in some infants with HIV-related illness,[10-12] and asymptomatic HIV-positive infants[9] erode investigators' ability to determine the true rate of vertical transmission of HIV infection.

This study was initiated in an attempt to measure the impact of HIV-seropositive newborns on the hospital care system. Our strategy was to determine the proportion of HIV-seropositive newborns who are hospitalized for an AIDS-related diagnosis (ARD) during their first year of life and to characterize those hospitalizations. This information can be used to plan the inpatient services required for HIV-infected infants.

Methods

The Statewide Planning and Research Cooperative System (SPARCS), maintained by the New York State Department of Health, collects diagnostic, demographic, and financial data on every hospital discharge in New York State.[13] Data on each hospital discharge include patient demographic information and primary and secondary diagnostic and procedure codes.

All hospital discharges from December 1987 through June 1989 were extracted from the SPARCS data base for any infant with a date of birth from December 1987 through June 1988 and an ARD of ICD-9CM 042, 043, 044, 795.8, or other selected codes (see Appendix) in any of the five diagnostic code categories. If a hospitalization had more than two ARDs, the code with the highest designation in the hierarchy listed above was used in the analysis.

Since SPARCS records represent hospital discharges and not individuals, multiple hospital discharges of an individual were collapsed into a single longitudinal case by linkages based on the date of birth, gender, and county of residence. Additional data elements such as dates of admission and discharge, medical record number, zip code of residence, race/ethnicity, patient disposition, and transfer status were examined to validate the case formation.

The number of HIV-seropositive infants who were born in each month from December 1987 to June 1988 was determined from the Newborn HIV Seroprevalence Study.[14]

The newborn seroprevalence and SPARCS data bases do not contain personal identifiers and cannot be linked. The analysis was therefore restricted to comparing, for each month of birth, the number of HIV-seropositive newborns and the number of infants hospitalized with an ARD within the first year of life. Frequencies were tabulated on the number of hospitalizations, the length of stay, each AIDS-related ICD-9-CM diagnostic code, age at first hospitalization, and case fatality rates.

Results

A total of 995 HIV-seropositive newborns born December 1987 through June 1988 were confirmed in the Newborn

TABLE VIII-1—Ratio of Infants Hospitalized for ARD within 12 Months after Birth to HIV-Seropositive Newborns, by Month of Birth

Month of Birth	ARD Cases	Number of HIV+ Newborns	Percent of ARD Cases/HIV+ Newborns
December 1987	26	169	15
January 1988	19	139	14
February 1988	18	122	15
March 1988	19	132	14
April 1988	24	135	18
May 1988	22	136	16
June 1988	23	162	14
Total	151	995	15

TABLE VIII-2—Hospitalizations within 12 Months of Birth by ARD at First Hospitalization

	Number of Hospitalizations		
	One	Two or More	
ARD at First Hospitalization	No. (Percent)	No. (Percent)	Total N
042	35 (73)	13 (27)	48
043/044	22 (79)	6 (21)	28
795.8	28 (64)	16 (36)	44
Other	25 (81)	6 (19)	31
Total	110 (73)	41 (27)	151

TABLE VIII-3—Deaths and Case Fatality Rates within 12 Months of Birth by Number of Hospitalizations

Number of Hospitalizations	No. of Deaths	No. of Cases	% Case Fatality Rate
One	23	110	21
Two +	8	41	20
Total	31	151	21

bers of the hospitalized set were members of the seroprevalence set, 15 percent of HIV-seropositive infants born in these seven months were hospitalized with an ARD within 12 months of birth. The monthly ratios ranged from 14 to 18 percent (Table VIII-1).

The 151 ARD cases had a total of 240 hospitalizations during the 12 months of follow-up. Most (73 percent) had only one hospitalization, 16 percent had two, 5 percent had three, and 6 percent had four to eight. Cases with a diagnostic code of 795.8 had the greatest proportion of multiple admissions (Table VIII-2).

The multiple admission patterns are influenced by a high case fatality rate of 21 percent (Table VIII-3). Of the 31 deaths, 23 (74 percent) occurred during the first hospitalization. As would be expected, cases with a 042 (AIDS) diagnostic code had the highest fatality rate (40 percent), compared with 18 percent for those with a 043 or a 044 code and 9 percent for those with a 795.8 or other code.

The 240 hospitalizations of the 151 cases comprised a total of 5,772 patients days, ranging from one to 368 days per hospitalization. Nearly half of the hospitalizations had a length of stay (LOS) exceeding 10 days (Figure 1). Cases with 042 codes had the greatest median LOS, 15 days (Table VIII-4). The other four categories had similar median LOS, ranging from 7.5 to 9.0 days.

Twenty-eight percent of the cases were first hospitalized for an ARD during the birth stay or within 90 days of birth (Figure 2). More than half (56 percent) were first hospitalized within six months of birth. Fifteen cases had an ARD at the hospitalization of birth. Seven of these 15 had a diagnostic code of 795.8; all 15 had an accompanying illness, such as severe infection or drug withdrawal. The average LOS for an ARD at the hospitalization of birth was 68.6 days.

Discussion

In this study, 151 of 995 (15 percent) newborns positive for HIV antibody were hospitalized with an ARD during the first year of life. Assuming maternal transmission rates of 25 to 50 percent,[3–10,15,16] it is estimated that of the 995 HIV-positive newborns, 250–498 infants were actually HIV-infected. Thus approximately 30–60 percent of the HIV-infected infants were hospitalized with an ARD within the first year of life.

The accuracy of this estimate is difficult to assess due to problems inherent in ecologic comparisons, including migration, deaths occurring outside the hospital, underreporting or inaccurate reporting of ARD cases to SPARCS, and missing and unknown data in both data bases. To minimize the impact of migration, we only included New York State residents from both files. However, 6 percent of the ARD cases included in this study had a county of residence at first hospitalization that was not a county of residence at birth for HIV-seropositive infants born in the same month.

Undercounting of ARD cases could occur when an infant had HIV-related illness, such as a bacterial infection, but was not coded as having HIV infection. This type of underreporting was assessed among 199 adult AIDS cases identified through a one-day hospital bed census and verified through chart review. SPARCS records were identified for 188 of those cases, and 177 indicated one of the current or discontinued HIV codes (personal communication, unpublished report, Benedict I. Truman, MD, Gerald Kaufman, PhD). To assess the accuracy of pediatric AIDS reporting to SPARCS, the AIDS Surveillance Program conducted chart reviews on 56 infants reported to SPARCS with a 042 diagnostic code in 1986. Half the cases were CDC-defined AIDS, 40 percent were AIDS-related conditions or suspect AIDS, and 10 percent were either adult or duplicate cases (personal communication, unpublished report, Benedict I. Truman, MD, Gerald Kaufman, PhD).

These analyses suggest that SPARCS data capture approximately 90 percent of HIV-related hospitalizations but are not very accurate in distinguishing AIDS from

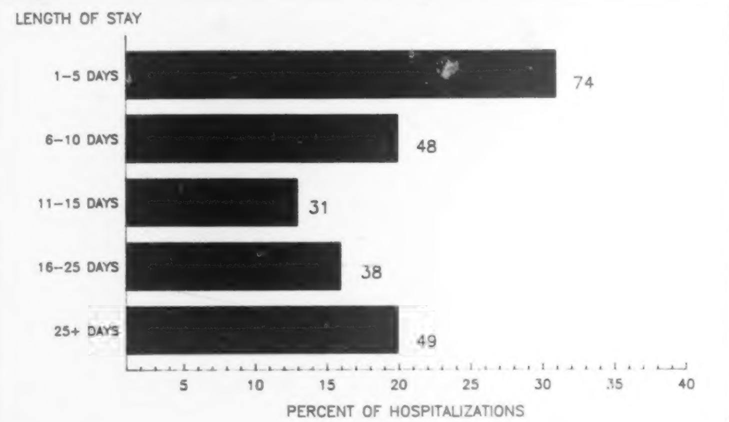

FIGURE 1—Distribution of Hospitalizations of ARD Cases by Length of Stay (in Days)

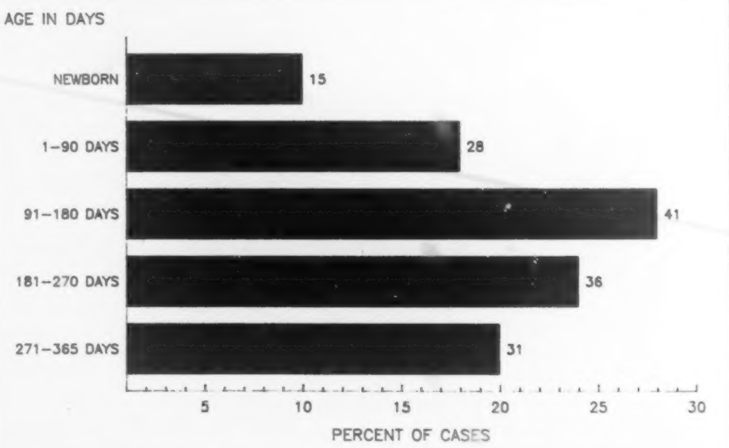

FIGURE 2—Distribution of ARD Cases by Age in Days at First Hospitalization

ARD. The latter finding was also substantiated in this study, in which nine cases switched from a 042 diagnosis to a 043, 044, or 795.8 diagnosis in subsequent hospitalizations.

Our finding that 15 percent of seropositive newborns are hospitalized with an ARD within 12 months of birth is consistent with two recent reports on pediatric AIDS.[9,10] The first study, conducted in Zaire, included HIV-seropositive infants who had lived beyond one month of age. Of the 412 infants studied, 13.4 percent had developed AIDS within 12 months of life, according to a definition based on the WHO (World Health Organization) clinical case definition of pediatric AIDS in Africa.[10] In the second study, from France, 19 (11 percent) of 177 seropositive infants developed severe forms of HIV infection within 18 months, based on the CDC classification for children.[9]

These two studies are not directly comparable with ours, because they used different study populations and case definitions. The French and Zaire studies were facility-based, whereas ours used the population-based Newborn HIV Seroprevalence Study and hospital discharge data. In addition, due to confidentiality concerns, our study could not identify HIV-seropositive or ARD cases but relied on an ecologic comparison based on month of birth. However, the similarity of results in the three studies, along with the internal consistency of month-to-month ARD/HIV ratios in this study, lend support to the accuracy of these estimates.

The results of this study are based on seven months (December 1987 through June 1988) of newborn seroprevalence data. This procedure will be conducted on an ongoing basis to project future hospital care needs of HIV-infected infants and to provide more information regarding maternal transmission and latency of HIV-related diseases in this population.

Summary

The purpose of this study was to estimate the proportion of HIV-seropositive newborns hospitalized for an AIDS-related diagnosis within 12 months of birth and to characterize these hospitalizations. The number of HIV-seropositive infants born December 1987 through June 1988 was derived from the blinded New York State Newborn HIV Seroprevalence Study. The number of infants who were born in these same months and hospitalized with an ARD within 12 months of birth was obtained from a data base containing information on all hospital discharges in New York State. Comparisons were made on the basis of month of birth.

Of infants born during the study period, 995 tested positive for HIV antibody, and 151 infants (15 percent) were hospitalized with an ARD within 12 months of birth. The 151 ARD cases had a total of 240 hospitalizations, accounting for 5,772 patient days. Fifty-six percent of the ARD cases were diagnosed within six months of birth. ☐

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TABLE VIII-4—Length of Stay for ARD Hospitalizations by Diagnosis

Diagnosis (ICD-9CM Code)	No. of Hospitalizations	Length of Stay (in Days)			
		Range	Median	Mean	SD
042	77	1-368	15.0	27.1	47.89
043	12	2-163	7.5	36.0	55.19
044	43	1-227	9.0	25.1	43.41
795.8	74	1-210	8.0	20.4	36.92
Other	34	2-205	7.5	19.4	41.23
Total	240	1-368	10.0	24.1	43.23

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APPENDIX

042	Human immunodeficiency virus infection with specified conditions
043	Human immunodeficiency virus infection causing other specified conditions
044	Other human immunodeficiency virus infection
795.8	Positive serological or viral culture findings for HIV

"Other" AIDS-Related Diagnosis

279.19	Deficiency of cell-mediated immunity
279.3	Unspecified immune deficiency
136.3	*Pneumocystis Carinii Pneumonia* (PCP)
173	Other malignant neoplasm of skin (for Kaposi's Sarcoma, KS)
117.5	Cryptococcosis
321.0	Cryptococcal meningitis
031	Mycobacterial Disease (non-TB)
130.0	Toxo meningo encephalitis
54.7	Herpes simplex w/other complications
200.0	Reticulosarcoma
200.8	Other named variants, lymphomas
516.8	Interstitial pneumonia
112.4	Candidiasis, lung
46.3	P.M. Leucoeneph
115	Histo

IX. HIV Seroprevalence in a Facility for Runaway and Homeless Adolescents

Rachel L. Stricof, MPH
James T. Kennedy, MD
Thomas C. Nattell, MA, MSW
Isaac B. Weisfuse, MD
Lloyd F. Novick, MD, MPH

Introduction

Despite the clinical and epidemiologic importance of HIV infection among adolescents, our knowledge of the extent of infection in this population in New York State has been limited, relying on AIDS case reports and data on military recruits. As of December 31, 1989, a total of 67 cases of AIDS in individuals 13–19 years old had been reported statewide, 45 males and 22 females. Their risk factors were reported as follows: 21 were homosexual/bisexual males, 12 were intravenous drug users, three were homosexual intravenous drug users, 10 were heterosexual contacts of persons at risk, 14 received blood products, three were foreign-born, and risk factors were not determined for four.

Military recruit data for New York State residents from October 1985 through December 1988 revealed a seroprevalence of 0.06 percent for persons 17–19 years old (communication from Centers for Disease Control; Quarterly Data Report on HIV in Civilian Applicants for Military Service, June 22, 1989). Thirty males and eight females were identified as seropositive for HIV.

Due to this limited availability of information on adolescents, a seroprevalence study was initiated in Covenant House New York, a facility serving runaway and homeless youth in New York City. Although this population is not representative of adolescents in general, it was selected as a sentinel group of high-risk youth. The study was designed to determine HIV seroprevalence among adolescents receiving health care at the facility, to ascertain selected demographic characteristics (age, race/ethnicity, gender), to assess trends in seroprevalence over time, and to determine the risk factors associated with HIV infection in this population.

Background

Covenant House New York predominantly serves adolescents under the age of 21. It consists of a crisis center, which can accommodate more than 300 individuals a night; a drug treatment program for 26 residents, with a 30-day average stay; a mother-child shelter; an outreach program (vans that go out at night); and a long-term transitional living program.

A recent survey[1] of 168 Covenant House adolescents revealed the following sociologic characteristics:

- The average length of time away from home was 90 weeks.
- Alcohol was used by 80 percent, marijuana by 68 percent, cocaine by 48 percent, and crack by 38 percent.
- Intravenous drug use was low (6 percent).
- Nearly all (91 percent) were sexually active, with an average of 2.8 sexual partners weekly (range, one to 20).
- Twenty-nine percent admitted to having exchanged sex for food, money, shelter, or drugs.
- Most (62 percent) reported having been arrested.
- Almost half (46 percent) reported having been treated as outpatients for psychiatric problems, and 18 percent had received inpatient psychiatric treatment.

Approximately 50–60 percent of the adolescents seeking shelter in Covenant House New York are seen in the health services department. Individuals are referred for health services immediately if they are on medication, placed in one of the group homes, a mother-infant pair, or deemed to have a health problem on intake. All others are referred if they remain in the facility for 24–48 hours. At the first visit a complete health assessment is performed, which includes a physical examination, urinalysis, complete blood count, and serologic test for syphilis.

TABLE IX-1—HIV Seroprevalence by Quarter among Runaway and Homeless Adolescents, October 1987–December 1989

Quarter	Number Positive	Number Tested	Percent Positive
4th 1987	22	280	7.9
1st 1988	14	315	4.4
2nd 1988	18	254	7.1
3rd 1988	20	292	6.9
4th 1988	11	250	4.4
1st 1989	16	300	5.3
2nd 1989	13	245	5.3
3rd 1989	20	415	4.8
4th 1989	8	316	2.5
Total	142	2,667	5.3

Methods

Study subjects were limited to Covenant House New York clients undergoing an initial medical evaluation between October 1, 1987 and December 31, 1989. Clients seen for annual or follow-up visits were excluded to avoid repeat testing of the same individuals over time. The study was limited to persons 15–20 years old, because few individuals under the age 15 or over age 20 visit the clinic and we wanted to be able to assure anonymity.

If a client met the inclusion criteria, an HIV seroprevalence study test requisition (SSTR) form was completed by the clinic staff. The completed form contained the adolescent's age, race or ethnicity, gender, and most recent county and zip code of residence, as well as the month and year of specimen collection. No personal identifier (e.g. name, previous street address, or medical record number) was recorded.

The SSTR form was held with the specimen until all routine laboratory tests were completed. If serum was left over, the laboratorian would pipette at least 0.5 ml into a tube labeled with the blind seroprevalence number and send the serum with the SSTR form to the New York State Department of Health's Wadsworth Center for Laboratories and Research for HIV testing.[2] No additional blood was drawn for this study. If there was not a sufficient quantity of blood left over, the demographic data alone were sent to the Department and entered into the data base.

In May 1988 a risk-assessment component was added to the study design. Information routinely obtained by the clinic staff and documented in the medical record was abstracted onto a separate risk assessment form with the blind seroprevalence number. This information included, where applicable, the gender of sex partners; the history of intravenous drug use (heroin or cocaine), crack use, prostitution, blood transfusion, and/or sexually transmitted disease; and the sex partners' history of drug use, homosexual/bisexual activity, or HIV status.

The information recorded on the SSTR form, including the HIV test result, was merged with the risk-assessment information and analyzed using computer software from SAS Institute, Inc; BMDP Statistical Software, Inc; the Centers for Disease Control[3]; and Epistat Services.[4]

Overall HIV seroprevalence rates were calculated by quarter-year of clinic visit. Chi-square tests for trend were applied to the data.[4]

Seroprevalence rates were calculated by gender, year of age, and for three racial/ethnic categories: non-Hispanic Whites, non-Hispanic Blacks, and Hispanics. The effects of the various demographic characteristics were assessed by multiple regression analysis after controlling for potential confounding effects of the others.

Univariate analysis for risk factors associated with HIV infection was performed with odds ratios and confidence intervals.[5] Mutually exclusive categories were established in the following hierarchical order: intravenous-drug-using homosexual/bisexual males, homosexual/bisexual males, intravenous drug users, blood transfusion recipients, persons with sex partners known or at risk for infection, persons with a history of sexually transmitted disease, prostitutes, and crack users. Individuals reporting none of these risk factors formed the reference group.

Demographic characteristics and/or risk factors independently associated with HIV infection were determined by multiple logistic regression after controlling for the effects of the others. Age, race/ethnicity, and risk factors derived from the univariate analysis were forced into the model. Adjusted odds ratios and confidence intervals were derived from the parameter estimates and standard errors.

Results

Over the 27-month study period, 2,667 specimens were analyzed for HIV antibody, and 142 (5.3 percent) were found to be positive. Five specimens were not tested due to insufficient serum, and eight equivocals were eliminated from the analysis.

The HIV seroprevalence rates by quarter are presented in Table IX-1. There was a tendency toward a decrease in seroprevalence over time, but the decline was not sufficient overall ($p = 0.27$) or for any subset of the population (males, females, non-Hispanic Whites, non-Hispanic Blacks, Hispanics, 15–16-year-olds, 17–18-year-olds, or 19–20-year-olds). A tendency toward a decline approaching statistical significance was found for females ($p = 0.06$), Blacks ($p = 0.08$), and 19–20-year-olds ($p = 0.09$).

Seroprevalence rates by age, race/ethnicity, and gender are presented in Table IX-2. The rates increased from 1.3 percent for 15-year-olds (not shown) to 8.6 percent for 20-year-olds. Hispanics had the highest seroprevalence rate, followed by non-Hispanic Whites and non-Hispanic Blacks. For males the rate was 6.0 percent; for females, 4.2 percent. However, after controlling for each of the other demographic characteristics, age was the only variable found to be associated with HIV seropositivity.

Between May 1988 and December 1989 risk-assessment forms were submitted on 1,758 individuals, 1,039 males and 719 females. The unadjusted HIV seroprevalence rates and univariate risk-factor analyses are presented in Table IX-3. HIV infection in males was associated with intravenous drug use, homosexual/bisexual activity, history of a sexually transmitted disease, prostitution, and crack use. HIV infection in females was associated with history of another sexually transmitted disease.

The results of the multivariate risk-assessment analyses are summarized in Table IX-4. After controlling for the confounding effects of the other variables listed, HIV infection in male adolescents was found to be associated with the following independent risk factors in declining order: homosexual/bisexual activity, prostitution, having had another sexually

TABLE IX-2—HIV Seroprevalence Rates by Demographic Characteristics among Runaway and Homeless Adolescents, October 1987–December 1989

Demographic Characteristics	Number Positive*	Number Tested*	Percent Positive	Odds Ratio**	95% Confidence Interval**
Age (years)					
15–16***	7	312	2.2	1.0	
17	10	340	2.9	1.1	0.4, 3.1
18	20	591	3.4	1.3	0.6, 3.2
19	49	750	6.5	2.7	1.2, 6.0
20	56	653	8.6	3.7	1.7, 8.3
Race/Ethnicity					
Non-Hispanic White	24	399	6.0	1.0	
Non-Hispanic Black	67	1,448	4.6	0.7	0.4, 1.2
Hispanic	47	688	6.8	1.1	0.7, 1.5
Gender					
Female	44	1,037	4.2	1.0	
Male	97	1,611	6.0	1.2	0.8, 1.8

*Differences in the total numbers found positive and the total numbers tested for each parameter reflect missing demographic data on the SSTR forms.

**Odds ratios and confidence intervals after controlling for the other demographic characteristics.

***Combined for logistic regression analysis due to small cell sizes.

TABLE IX-3—Univariate Analysis of Risk Factors for HIV Infection by Gender among Runaway and Homeless Adolescents, May 1988–September 1989

Hierarchical Risk Factor	Number Positive	Number Tested	Percent Positive	Odds Ratio	95% Confidence Interval
Males					
IVDU and Homosexual or Bisexual	4	12	33.3	47.5	8.0, 243.0
Homosexual or Bisexual	27	109	24.8	31.2	12.1, 94.7
IVDU	3	32	9.4	9.8	1.5, 48.4
Blood Transfusion	1	9	11.1	11.9	0.2, 116.6
Sex Partner at Risk	1	11	9.1	9.5	0.2, 89.9
STD	4	69	5.8	5.9	1.2, 25.3
Prostitution	2	7	28.6	38.0	2.9, 290.4
Crack Use	11	214	5.1	5.2	1.7, 17.1
No Risk Indicated	6	576	1.2	1.0	
Females					
IVDU	2	16	12.5	8.5	0.8, 47.9
Blood Transfusion	0	3	0.0	0.0	0.0, 158.7
Sex Partner at Risk	1	8	12.5	8.5	0.2, 79.9
STD	8	124	6.4	4.1	1.3, 12.8
Prostitution	2	16	12.5	8.5	0.8, 47.9
Crack Use	2	69	2.9	1.8	0.2, 9.1
No Risk Indicated	8	483	1.7	1.0	

Abbreviations: IVDU, intravenous drug use; STD, sexually transmitted disease (other than HIV infection).

transmitted disease, and crack use. For females no risk factors were independently associated with HIV infection, although there was a tendency toward an association for prostitution and having had another sexually transmitted disease.

Discussion

This seroprevalence study of a sentinel group of runaway and homeless adolescents disclosed an alarming rate of HIV infection. More than 5 percent were seropositive. Although not representative of adolescents in general, they demonstrate the magnitude of the problem for a large and definable group at one end of the spectrum. These findings have implications, not only for clients of this facility, but also for other groups who interact with runaways and homeless and are consequently at risk.

These homeless youths come from the streets of New York City; their environment is characterized by drug use, prostitution, poverty, and crime. The seroprevalence data indicate an association between HIV infection and prostitution, history of other sexually transmitted diseases, and crack use, in addition to other, previously defined risk factors. Exchanging sex for food, money, shelter, or drugs with multiple sexual partners, some of whom are infected, is the most likely explanation for HIV transmission in this group of adolescents.

Since we were limited to obtaining risk information that was documented in the medical record during the standard course of medical practice, the data probably do not reflect the full incidence of these risk behaviors and should be interpreted with caution. Data provided by Margetson and Lipman[1] suggest that prostitution and crack use were underreported in our study.

The adolescents in this study are highly susceptible hosts living in a hazardous environment. They are regularly exposed to infectious agents, including HIV. Host factors include high levels of sexual activity, unprotected sexual intercourse, and homosexual/bisexual activity in males. Environmental factors include poverty and the necessity to support oneself and obtain drugs in an environment where drugs, prostitution, sex, and infection are commonplace and linked. Prevention activities face the challenge of inducing change, not only in the individuals, but in the external factors that facilitate exposure.

Prevention initiatives must start earlier than the adolescent period if a satisfactory impact is to be attained. All schools in New York State must provide basic HIV education. Regional training coordinators from the New York State Education Department are helping to develop and implement these programs.

The New York State Department of Health, using state and federal funds, has contracted with a variety of community-based organizations to provide HIV education, counseling, and testing services as well as special living arrangements or referrals, basic needs (food, clothing, shelter, etc.), psychosocial support, and medical services to high-risk adolescents.

Summary

In October 1987, the New York State Department of Health initiated a study to determine the prevalence of antibody to HIV in clients of a facility for runaway and homeless adolescents in New York City. A risk-assessment component was added in May 1988. As of December 1989, a total of

TABLE IX-4—Multivariate Analysis of Factors Associated with HIV Infection among Runaway and Homeless Adolescents, May 1988–December 1989

Demographic Characteristic or Risk Factor	Odds Ratio (95% Confidence Interval)*	
	Males	Females
Race/Ethnicity		
Non-Hispanic White	1.0	1.0
Non-Hispanic Black	1.0 (0.4, 2.6)	1.1 (0.3, 3.8)
Hispanic	2.0 (0.8, 5.2)	0.6 (0.1, 3.3)
Homosexual/Bisexual Male	6.1 (2.8, 12.9)	—
Intravenous Drug Use	0.9 (0.3, 2.9)	1.9 (0.3, 11.7)
Crack Use	2.3 (1.2, 4.3)	1.3 (0.4, 3.9)
History of STD**	2.4 (1.1, 5.0)	2.3 (0.9, 6.2)
Prostitution	2.8 (1.0, 7.5)	4.0 (0.9, 18.3)

*After controlling for age.
**Sexually transmitted disease (other than HIV infection).

2,667 adolescents had been tested, and 142 (5.3 percent) were found to be HIV-seropositive (males 6.0 percent, females 4.2 percent). The seroprevalence rate increased from 1.3 percent for 15-year-olds to 8.6 percent for 20-year-olds. Hispanics had the highest seroprevalence rate (6.8 percent), followed by non-Hispanic Whites (6.0 percent) and non-Hispanic Blacks (4.6 percent). HIV seropositivity was associated with intravenous drug use, male homosexual/bisexual activity, prostitution, and history of another sexually transmitted disease. The alarmingly high prevalence of HIV infection in this selected population illustrates the immediate need for prevention programs for adolescents. □

Acknowledgments

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X. The AIDS Epidemic in New York State

Perry F. Smith, MD
Jaromir Mikl, MSPH
Sue Hyde, BS
Dale L. Morse, MD, MS

Introduction

Surveillance for cases of acquired immunodeficiency syndrome (AIDS) has provided important information for characterizing and controlling the epidemic. Even before the human immunodeficiency virus (HIV) had been identified in 1983 and 1984,[1,2] AIDS surveillance demonstrated that gay men and intravenous drug users (IVDUs) bore a disproportionate share of the burden of disease.[3] Surveillance data provided the first information about the modes of transmission of HIV and were thus the basis for early recommendations to limit its spread.[4]

New York State reported some of the first AIDS cases and has the highest number of cases in the nation, mostly from New York City. This report describes the AIDS epidemic statewide and compares it with the epidemic in the rest of the nation. It also compares data from the state's Newborn HIV Seroprevalence Study with AIDS surveillance data for women.

Methods

Both the New York State Department of Health (DOH) and New York City DOH maintain active surveillance for AIDS cases, including regular solicitation of reports from hospitals. New York City DOH staff investigate all cases reported within the city; New York State DOH staff investigate cases in New York State exclusive of New York City. Once a suspect case is identified, trained surveillance staff review the patient's medical records to collect demographic, risk-factor, and clinical information. Vital status is regularly updated by systematic review of death certificate reports and by encouraging hospitals to report deaths to surveillance staff.

Throughout the epidemic, all suspect cases have been evaluated according to the national surveillance case definition, which has been expanded twice, in 1985 and 1987.[3,5,6] Each case was evaluated by the criteria current when the case was investigated. Cases meeting more than one definition were classified hierarchically according to the earliest definition that they met.

Special case ascertainment studies indicate that 75-85 percent of diagnosed AIDS cases in New York City are eventually reported,[7] as are more than 95 percent of patients discharged with AIDS-related diagnoses in New York State exclusive of New York City.[8]

Both the State and New York City DOH report their confirmed cases independently, without identifying information, to the Centers for Disease Control (CDC) for national surveillance. The New York State DOH also maintains a statewide computerized AIDS registry in which the state and city case files are merged.

For comparison with the statewide case data, the number of cases in the rest of the United States was determined from summary national AIDS surveillance data reported to CDC through November 1989.[9] For each subgroup comparison the number of cases reported statewide was subtracted from the corresponding national total. To ensure comparability between the statewide and national data, we selected all cases registered in New York State exclusive of New York City as of October 31, 1989, and all cases registered in New York City as of November 17, 1989—the final dates of data transfer for the national surveillance report.

Geographic classification of persons with AIDS was based on their residence at the time they met the national surveillance case definition, except for cases outside New York City before July 1989, where we used the residence at onset of symptoms. Cases were also classified as adults/

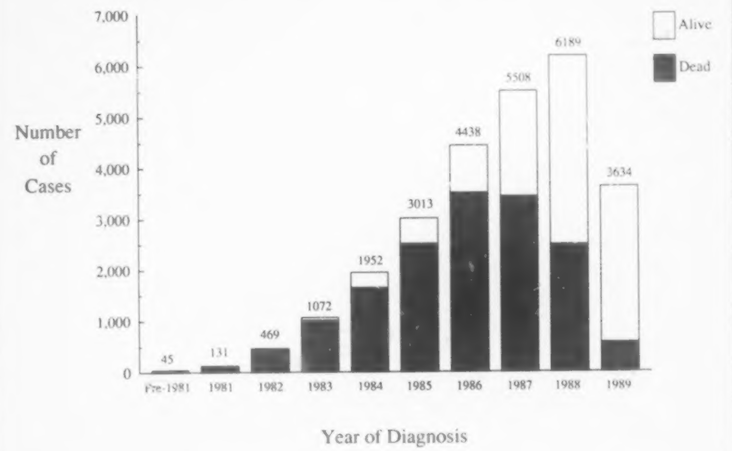

FIGURE 1—Number of AIDS cases (N = 26,576) by Vital Status and Year of Diagnosis, New York State

NOTE: A total of 125 cases with unknown vital status were excluded. Data for 1989 are incomplete due to the delay in case reporting.

adolescents (>12 years old) or children (<13 years old) and by hierarchical exposure groups, according to the CDC system.[9]

As a measure of HIV transmission among adolescents, data for persons 20–24 years old were analyzed. Adults in this age group are likely to have been infected during adolescence because of the long median incubation period for AIDS, which studies suggest is about 10 years.[10–12] Cases among teenagers were not included in this analysis, since their infections may have occurred before adolescence.

AIDS incidence rates were calculated with official New York State population projections by age, sex, and geographic region for 1985.[13] For race-specific population estimates, we applied the racial distribution from the 1980 Census to the 1985 projections, since race-specific projections are not available.

For comparison with results of the Newborn HIV Seroprevalence Study,[14] we used the incidence of AIDS per 100,000 women of childbearing age (15–49 years) during 1988. From the Newborn HIV Seroprevalence Study, we used data for all births to mothers in this age group in 1988. The data were stratified by age, race/ethnicity, and geographic regions, according to urbanization and proximity to New York City, as described elsewhere.[14]

Results

Cases by Type of Definition

A total of 26,576 cases of AIDS were reported statewide by the cutoff dates for this report, accounting for 23 percent of the 115,158 cases reported nationally. Of these cases, 22,186 (83 percent) met the pre-1985 national surveillance case definition,[3] 592 (2 percent) met the 1985 definition,[5] and 3,798 (14 percent) met the expanded 1987 definition.[6] Of the 1987 definition cases, 47 percent had indicator diseases that were diagnosed definitively; the rest were diagnosed presumptively. Of 9,847 cases diagnosed in 1988 and 1989, 29 percent met only the 1987 case definition (13 percent definitive, 16 percent presumptive). Of the total 26,576 cases, 16,163 (61 percent) had *Pneumocystis carinii* pneumonia, and 3,505 (13 percent) had Kaposi's sarcoma.

Incidence and Mortality

The number of AIDS cases diagnosed statewide has increased every year through 1988 (Figure 1). An estimated 90 percent of cases that are ultimately reported are included in the registry by 11 months after diagnosis (Dr. Lawrence Lessner, New York State Department of Health, personal communication). Because of this delay in reporting, the number of cases for 1989 is expected to exceed that for 1988 when reporting is complete.[15]

Of the 26,576 reported cases, 17,531 (66 percent) are known to have died as of the cutoff dates for this study. Thirteen percent of all cases died within one month of diagnosis. The case fatality rate is highest for cases with the longest follow-up: 90 percent of cases diagnosed before 1985. Of patients diagnosed before 1988, 80 percent have died (median survival time, 10.3 months).

Geographic Distribution

New York City accounts for 22,231 (84 percent) of all statewide cases, although the percentage has dropped as the epidemic spread. Another 3,250 cases (12 percent) were reported among the general population from New York State exclusive of New York City and 1,095 (4 percent) among prisoners in the state correctional system.

The geographic distribution of cases in quartiles is shown in Figure 2. Cases among prisoners are excluded, because accurate information on their counties of residence was not available. Although cases are widely distributed throughout the state, they tend to concentrate in counties with major metropolitan areas (e.g., Albany, Erie, Monroe, and Onondaga), as well as in New York City and nearby counties.

Gender

Males account for 22,482 (87 percent) of the statewide adult/adolescent cases (Table X-1). There is no difference in gender distribution between cases in New York City and New York State exclusive of New York City. The male case predominance reflects the large number of cases among men with a history of homosexual/bisexual contact. However, this percentage is lower than that for the rest of the United States, because New York State has a disproportionately high number of women who contracted AIDS by intravenous drug use.

Among children there is a slightly higher proportion of cases among males, but the gender distribution differs little from the rest of the United States (Table X-2).

Age

The age distribution of adult/adolescent cases statewide is similar to that in the rest of the United States, with nearly 90 percent of cases occurring among persons 20–49 years old (Table X-1).

Among children, New York State has 604 reported cases statewide and accounts for 31 percent of all such cases in the United States (Table X-2). Of these pediatric patients, 534 (88 percent) were

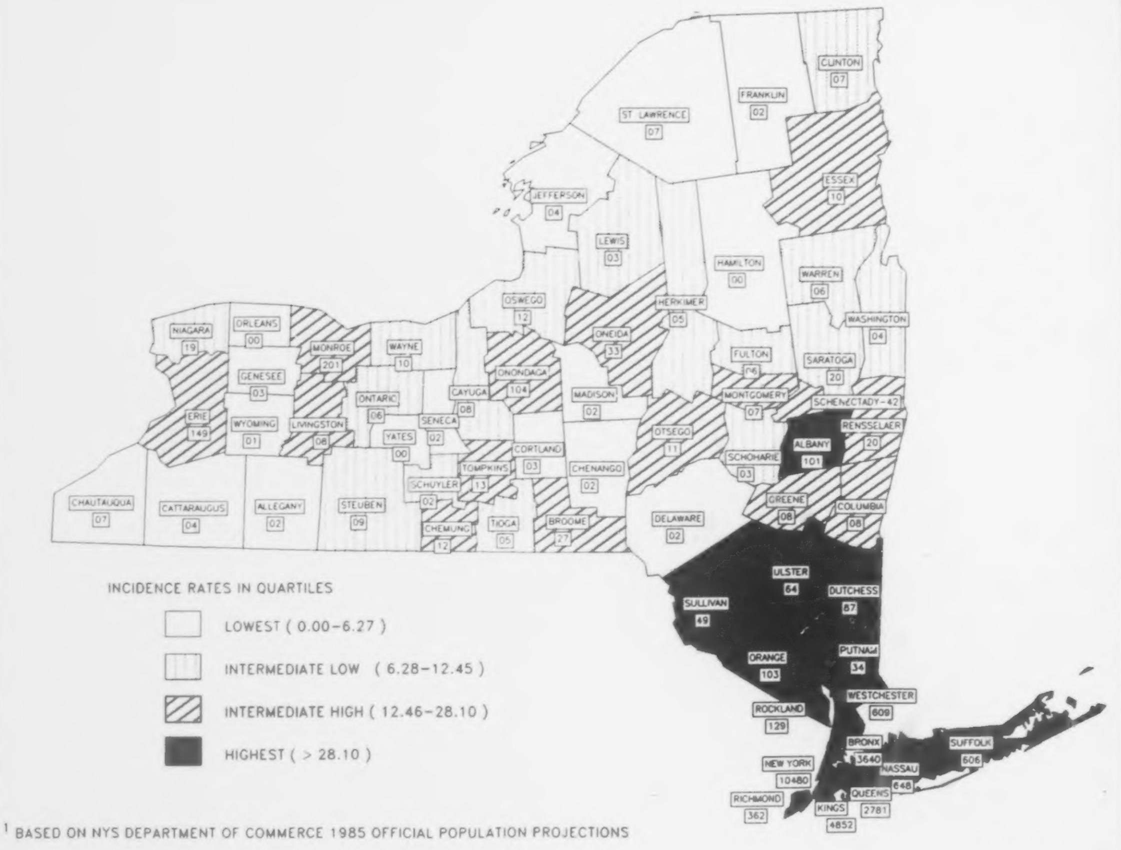
CLINTON 07
FRANKLIN 02
ST. LAWRENCE 07
ESSEX 10
JEFFERSON 04
LEWIS 03
HAMILTON 00
WARREN 06
OSWEGO 12
HERKIMER 05
WASHINGTON 04
NIAGARA 19
ORLEANS 00
MONROE 201
WAYNE 10
ONEIDA 33
FULTON 06
SARATOGA 20
GENESEE 03
ONONDAGA 104
ONTARIO 06
CAYUGA 08
MADISON 02
MONTGOMERY 07
SCHENECTADY - 42
ERIE 149
WYOMING 01
LIVINGSTON 08
SENECA 02
YATES 00
CORTLAND 03
OTSEGO 11
ALBANY 101
RENSSELAER 20
TOMPKINS 13
SCHUYLER 02
CHENANGO 02
SCHOHARIE 03
GREENE 08
COLUMBIA 08
CHAUTAUQUA 07
CATTARAUGUS 04
ALLEGANY 02
STEUBEN 09
CHEMUNG 12
TIOGA 05
BROOME 27
DELAWARE 02
ULSTER 64
DUTCHESS 87
SULLIVAN 49
ORANGE 103
PUTNAM 34
WESTCHESTER 609
ROCKLAND 129
BRONX 3640
SUFFOLK 606
NEW YORK 10480
NASSAU 648
QUEENS 2781
RICHMOND 362
KINGS 4852
INCIDENCE RATES IN QUARTILES
LOWEST (0.00-6.27)
INTERMEDIATE LOW (6.28-12.45)
INTERMEDIATE HIGH (12.46-28.10)
HIGHEST (> 28.10)
1 BASED ON NYS DEPARTMENT OF COMMERCE 1985 OFFICIAL POPULATION PROJECTIONS

FIGURE 2—Cumulative number of AIDS Cases and Incidence per 100,000 Population since the beginning of the Epidemic, by County or Borough of Residence, New York State

NOTE: Cases among New York State prison inmates were excluded. Cumulative incidence was based on New York State Department of Commerce population projections for 1985.

residents of New York City. Compared with the rest of the United States a greater proportion of cases occurred among children under five years old, reflecting the greater proportion of children infected perinatally in this state.

Race/Ethnicity

Whereas Blacks and Hispanics account for 37 percent of adult/adolescent cases in the rest of the United States, they account for nearly 60 percent of such cases statewide (Table X-1). This difference reflects the large percentage of Black and Hispanic cases in New York City (62 percent) and in state prisons (85 percent). In New York State exclusive of New York City, Blacks and Hispanics account for 37 percent of cases.

The statewide cumulative incidence rate (per 100,000 population) from the beginning of the epidemic among adult/adolescent males is several times higher among Blacks (901) and Hispanics (1,029) than among Whites (183) (Table X-3). For females, the racial/ethnic disparity is even greater (Table X-4).

Of children with AIDS, nearly 90 percent statewide are Black or Hispanic, compared to 72 percent in the rest of the United States (Table X-2).

Exposure Groups

The AIDS epidemic among adults and adolescents has primarily affected two exposure groups: men with a history of homosexual/bisexual contact, and IVDUs. The epidemic curve for gay men has leveled off or declined in recent years (Figure 3).

In 1988, for the first time, more cases were diagnosed among IVDUs than among gay men statewide. The percentage of cases among IVDUs is greater statewide (39 percent) than in the rest of the United States (16 percent) (Table X-1). Of the 1,095 prisoners with AIDS, 957 (87 percent) gave a history of intravenous drug use without homosexual/bisexual contact. Even when prisoners are excluded, the percentage of cases with a history of intravenous drug use is still higher in New York City (36 percent) and in New York State exclusive of New York City (35 percent) than in the rest of the United States (16 percent). The percentage of

TABLE X-1—Distribution of AIDS Cases among Adults/Adolescents (Age 13 and Older) in New York State and in the Rest of the United States

	New York State		United States (Excluding New York State)	
Demographic Variables	Number	(%)	Number	(%)
Sex				
Males	22,482	(86.6)	80,360	(92.0)
Females	3,490	(13.4)	6,879	(8.0)
Age Group (years)				
13–19	65	(0.3)	382	(0.4)
20–29	4,568	(17.6)	18,930	(21.7)
30–39	12,737	(49.0)	40,437	(46.4)
40–49	6,010	(23.1)	18,337	(21.0)
50+	2,592	(10.0)	9,157	(10.5)
Race/Ethnicity				
White	10,185	(39.2)	54,008	(61.9)
Black	8,729	(33.6)	21,758	(24.9)
Hispanic	6,753	(26.0)	10,655	(12.2)
Other	163	(0.6)	685	(0.8)
Unknown	142	(0.5)	133	(0.2)
Exposure Category				
Male Homosexual/Bisexual Contact	11,995	(46.2)	56,572	(64.8)
Intravenous (IV) Drug Use	10,120	(39.0)	13,602	(15.6)
Male Homosexual/Bisexual Contact and IV Drug Use	1,137	(4.4)	6,786	(7.8)
Hemophilia/Coagulation Disorder	61	(0.2)	986	(1.1)
Heterosexual Contact	1,390	(5.4)	4,067	(4.7)
Sex with Person at Risk for AIDS	903	(3.5)	2,951	(3.4)
Born in Pattern II Country*	487	(1.9)	1,116	(1.3)
Receipt of Blood Transfusion, Blood Components, or Tissue	267	(1.0)	2,501	(2.9)
Other/Undetermined	1,002	(3.9)	2,725	(3.1)

*Includes persons presumed to have become HIV-infected through heterosexual contact because they were born in countries where heterosexual transmission is believed to be the major mode of transmission.[9]

cases among non-IVDU men with a history of homosexual/bisexual contact was similar in New York City (49 percent) and New York State exclusive of New York City (44 percent).

The third most common exposure group is persons at risk for AIDS because of heterosexual contact. A slightly higher percentage of cases is in this category statewide than in the rest of the United States (5.4 vs 4.7 percent) (Table X-1). This category contains two subgroups: persons who report heterosexual contact with a person at risk for AIDS, and persons who are presumed to have become HIV-infected through heterosexual contact because they were born in countries where heterosexual transmission is believed to be the major mode of transmission, the so-called Pattern II countries.[9] Of the 1,390 statewide cases of heterosexual contact, 903 (65 percent) reported contact, and 487 (35 percent) were from Pattern II countries. Of the 903 cases with reported contact, 757 (84 percent) reported having sex with an IVDU.

The distribution of exposure categories varied by race/ethnicity and gender. Of all cases among White males, 78 percent were attributed solely to homosexual/bisexual contact, in contrast to 35 percent of Black and Hispanic males (Table X-3). Intravenous drug use was the leading exposure category for both Black (49 percent) and Hispanic (55 percent) males. Among females there was less variation, although heterosexual contact accounted for a larger percentage of cases among Blacks (27 percent) and Hispanics (34 percent) than among Whites (23 percent) (Table X-4). Overall the leading exposure categories for females were intravenous drug use (62 percent) and heterosexual contact with a person at risk for AIDS (25 percent). Of the 864 females in the latter group, 729 (84 percent) reported having sex with an IVDU.

Statewide, 91 percent of children with AIDS were classified as acquiring their infection from their mothers (perinatally), compared with 76 percent in the rest of the United States (Table X-2). This difference is largely due to the majority (59 percent) of pediatric cases statewide whose mothers used intravenous drugs.

AIDS among Young Adults

Statewide 867 AIDS cases were reported among persons 20–24 years old, of whom 251 (29 percent) were White, 298 (34 percent) were Black, and 309 (36 percent) were Hispanic; the remaining nine were of other or unknown race/ethnicity. The leading exposure groups for these cases were male homosexual/bisexual contact, 403 (46 percent); intravenous drug use, 248 (29 percent); male homosexual/bisexual contact and intravenous drug use, 40 (5 percent); heterosexual contact with an intravenous drug user, 64 (7 percent); and born in a Pattern II country, 29 (3 percent). The remaining 83 cases were attributed to other exposure categories.

Comparison of AIDS Data with Newborn HIV Seroprevalence

Two measures of HIV infection among women are shown in Table X-5: the incidence of AIDS per 100,000 women in 1988, and the rate of HIV seropositivity among women giving birth in 1988, based on data from the Newborn HIV Seroprevalence Study. The absolute rates of these measures varied because of the different events they reflect: onset of severe HIV disease and prevalence of infection among women giving birth. However, both measures showed comparable trends. The highest rates occurred among persons 20–39 years old, Blacks, Hispanics, and women residing in New York City. Both rates declined with distance from New York City. The only disproportionate trend was among persons 15–19 years old: a disproportionately lower rate of AIDS per 100,000 women relative to the older groups, compared with the HIV seropositivity rates from the newborn study.

Discussion

AIDS has become a major cause of morbidity and mortality statewide. It is the leading cause of death among men 30–49 years old and women 20–39 years old in New York City, as well as among persons 30–39 years old in minority racial/ethnic groups in New York State exclusive of New York City.[15]

Statewide the effect of the AIDS epidemic among IVDUs is apparent in its increasing toll on racial minorities, women, and children. Transmission of HIV by contaminated needles has disproportionately affected Blacks and Hispanics. Both intravenous drug use and heterosexual transmission from IVDUs have caused a greater proportion of cases among women here than in the rest of the nation, and these women have transmitted HIV perinatally to a large number of children.

AIDS surveillance data have complemented the information provided by the HIV seroprevalence studies. Whereas seroprevalence studies provide timely information on the number, geographic distribution, and changing characteristics of HIV infected persons, information from AIDS surveillance provides detailed information on each reported case since the beginning of the epidemic. AIDS surveillance data show patterns similar to those from the Newborn HIV Seroprevalence Study except for one age-group difference. Where the Newborn Seroprevalence Study shows an alarming rate of HIV infection among women 15–19 years old, surveillance data show that they have a low incidence of AIDS. This finding appears to reflect the long incubation period for AIDS. AIDS case data for persons 20–24 years old statewide suggest that HIV transmission among adolescents has occurred more frequently among Blacks and Hispanics than among Whites and that sexual contact (both homosexual and heterosexual) has been the leading risk behavior.

TABLE X-2—Distribution of AIDS Cases among Children (Age 12 and Under) in New York State and in the Rest of the United States

Demographic Variables	New York State Number	New York State (%)	United States (Excluding New York State) Number	United States (Excluding New York State) (%)
Sex				
Males	320	(53.0)	740	(55.1)
Females	284	(47.0)	603	(44.9)
Age Group (years)				
<5	531	(87.9)	1,071	(79.7)
5–12	73	(12.1)	272	(20.3)
Race/Ethnicity				
White	62	(10.3)	363	(27.0)
Black	334	(55.3)	693	(51.6)
Hispanic	205	(33.9)	272	(20.3)
Other	1	(0.2)	12	(0.9)
Unknown	2	(0.3)	3	(0.2)
Exposure Category				
Mother at Risk for AIDS	550	(91.1)	1,022	(76.1)
Intravenous (IV) Drug Use	359	(59.4)	449	(33.4)
Sex with IV Drug User	112	(18.5)	210	(15.6)
Sex with Other Person at Risk for AIDS/HIV	17	(2.8)	104	(7.7)
Born in Pattern II Country*	47	(7.8)	120	(8.9)
Receipt of Blood Transfusion, Blood Components, or Tissue	3	(0.5)	30	(2.2)
Has HIV Infection, Risk Not Specified	12	(2.0)	109	(8.1)
Receipt of Blood Transfusion, Blood Components, or Tissue	19	(3.1)	189	(14.1)
Hemophilia/Coagulation Disorder	4	(0.7)	100	(7.5)
Undetermined	31	(5.1)	32	(2.4)

*See Table 1.

TABLE X-3—AIDS Cases among Male Adults/Adolescents by Exposure Category and Race/Ethnicity, and Cumulative Incidence per 100,000 Males, Age 15 Years and Older, in New York State

Exposure Category	White Number	White (%)	Black Number	Black (%)	Hispanic Number	Hispanic (%)	Total* Number	Total* (%)
Male Homosexual/Bisexual Contact	7,428	(77.6)	2,408	(34.8)	1,986	(34.8)	11,994	(53.4)
Intravenous (IV) Drug Use	1,365	(14.3)	3,418	(49.3)	3,129	(54.8)	7,965	(35.4)
Male Homosexual/Bisexual Contact and IV Drug Use	391	(4.1)	423	(6.1)	317	(5.6)	1,137	(5.1)
Hemophilia/Coagulation Disorder	45	(0.5)	4	(0.1)	7	(0.1)	56	(0.2)
Heterosexual Contact	17	(0.2)	375	(5.4)	4	(0.1)	398	(1.8)
Sex with Person at Risk for AIDS	17	(0.2)	17	(0.2)	4	(0.1)	39	(0.2)
Born in Pattern II Country**	0	(0)	358	(5.2)	0	(0)	359	(1.6)
Receipt of Blood Transfusion, Blood Components, or Tissue	119	(1.2)	25	(0.4)	20	(0.4)	168	(0.7)
Other/Undetermined	206	(2.2)	275	(4.0)	247	(4.3)	760	(3.4)
Total	9,571	(100.0)	6,928	(100.0)	5,710	(100.0)	22,478	(100.0)
Cumulative Incidence per 100,000***	183.1		901.1		1,028.9		336.2	

*Included in the total are 149 cases whose race/ethnicity is American Indian, Alaskan Native, Asian, or Pacific Islander and 120 whose race/ethnicity is unknown.

**See Table 1.

***Incidence was calculated as the cumulative number of AIDS cases among males, age 15 years and older, since the beginning of the epidemic, divided by estimates for the male population, age 15 years and older, by race/ethnicity in 1985. (See text for details.)

TABLE X-4—AIDS Cases among Female Adults/Adolescents by Exposure Category and Race/Ethnicity, and Cumulative Incidence per 100,000 Females, Aged 15 Years and Older, in New York State

	White		Black		Hispanic		Total*	
Exposure Category	Number	(%)	Number	(%)	Number	(%)	Number	(%)
Intravenous (IV) Drug Use	368	(60.2)	1,141	(63.4)	633	(60.7)	2,155	(61.8)
Hemophilia/Coagulation Disorder	2	(0.3)	1	(0.1)	0	(0)	3	(0.1)
Heterosexual Contact	142	(23.2)	488	(27.1)	352	(33.7)	992	(28.4)
Sex with Person at Risk for AIDS	142	(23.2)	360	(20.0)	352	(33.7)	864	(24.8)
Born in Pattern II Country**	0	(0)	128	(7.1)	0	(0)	128	(3.7)
Receipt of Blood Transfusion, Blood Components, or Tissue	56	(9.2)	27	(1.5)	12	(1.2)	97	(2.8)
Other/Undetermined	43	(7.0)	143	(7.9)	46	(4.4)	242	(6.9)
Total	611	(100.0)	1,800	(100.0)	1,043	(100.0)	3,489	(100.0)
Cumulative Incidence per 100,000***	10.5		185.5		160.0		46.0	

*Included in the total are 13 cases whose race/ethnicity is American Indian, Alaskan Native, Asian, or Pacific Islander and 22 whose race/ethnicity is unknown.

**See Table 1.

***Incidence was calculated as the cumulative number of AIDS cases among females, age 15 years and older, since the beginning of the epidemic, divided by estimates for the female population, age 15 years and older, by race/ethnicity in 1985. (See text for details.)

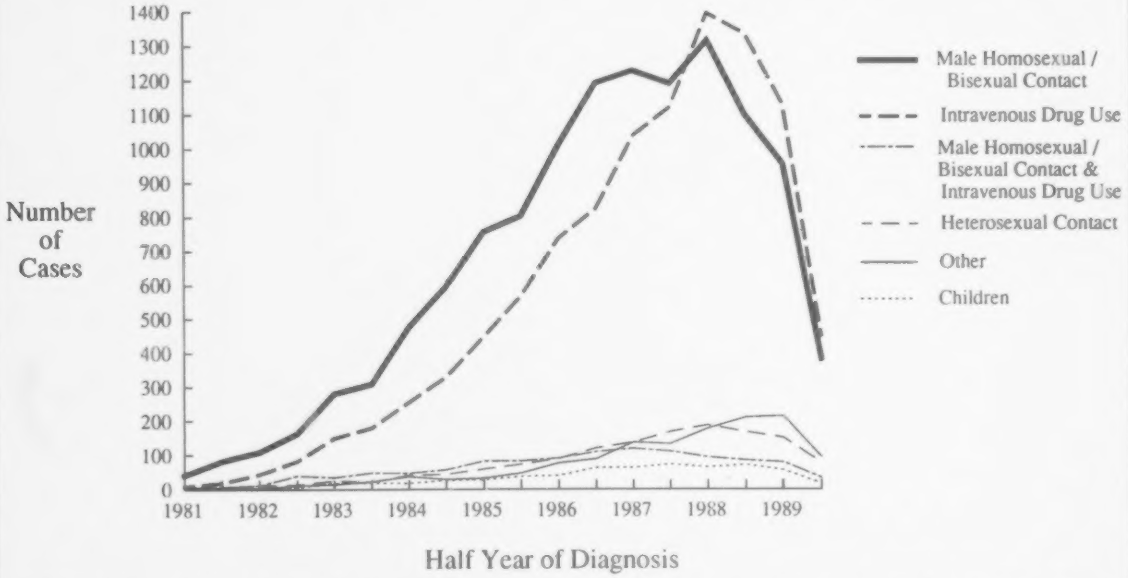

FIGURE 3—Number of AIDS cases among adult/adolescent exposure groups and among children by half-year of diagnosis, New York State

NOTE: Data for 1989 are incomplete due to the delay in case reporting.

The New York State DOH continues to monitor information from AIDS surveillance and the HIV seroprevalence studies to better address the needs of those already infected and to plan interventions that will limit the spread of HIV.

Summary

New York State has reported more AIDS cases (26,576) through November 1989 than any other state, accounting for 23 percent of all reported cases in the nation at the time. New York City has reported 22,231 (84 percent) of the statewide cases. The epidemiology of AIDS in New York State has been easily influenced by the large number of cases among IVDUs. Whereas intravenous drug use accounts for 16 percent of adult/adolescent cases in the rest of the nation,

TABLE X-5—Comparison of the Number and Incidence of Female AIDS Cases with HIV Seroprevalence among Women Giving Birth by Age Group, Race/Ethnicity, and Geographic Region in New York State 1988

Characteristics	Number of Cases and Incidence of AIDS per 100,000 Women in 1988 No.	Incidence*	Prevalence (per 100,000 Births) of HIV Seropositivity among Women Giving Birth in 1988** (N = 271,843)
Age Group (years)			
15–19	3	0.4	430
20–29	230	14.8	650
30–39	523	37.1	670
40–49	144	13.7	360
Race/Ethnicity			
White	145	4.0	120
Black	466	77.2	1,790
Hispanic	278	68.6	1,250
Other/Unknown	11	13.0	340
Geographic Region			
New York City	776	40.3	1,190
New York City Vicinity	80	7.9	260
Mid-Hudson Valley	13	5.8	220
Upstate Urban	24	3.5	130
Upstate Rural	7	0.8	80
Total	900	19.1	630

*Based on the number of AIDS cases among women ages 15–49 years and the 1985 population projections for women of this age group in New York State.

**Based on newborn HIV seroprevalence testing.

it accounts for 39 percent of the cases in New York State. Intravenous drug use is the leading exposure category among racial minorities, with a disproportionately high rate of cases among Blacks and Hispanics. The epidemic among women and children has also been severe, resulting from intravenous drug use by women, heterosexual transmission from drug users to women, and perinatal transmission to their children. ☐

Acknowledgments

The authors gratefully acknowledge the AIDS surveillance staff of the New York State and New York City Departments of Health for data collection, Rachel Stricof and Michael Davisson for newborn seroprevalence data, and Lloyd Novick for manuscript review. Surveillance activities were partially supported with funding through cooperative agreements with the Centers for Disease Control.

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rence in the community . . .[1] of a group of illnesses of similar nature, clearly in excess of normal expectation." The attribution of this particular epidemic to the adverse effect of postaccident stress provides new and potentially important insights into the etiology of cancer and into the circumstances that may eventually need to be managed to better protect the public health. The idea that stress may be an environmental culprit could change many aspects of the way we handle acute environmental problems like the Three Mile Island accident. The way such environmental events are publicized and managed may need to be reconsidered if further evidence supports the conclusion that stress can cause or promote a disease like cancer. Studies of cancer incidence at the Love Canal before the problem was publicized in 1978 provided no clear evidence of increased cancer risks even though chronic exposure to environmental toxins had taken place over years and decades.[3] However, since the entire neighborhood surrounding the Love Canal was relocated at the time the problem of toxic exposure was publicized in 1978, it has not been possible to determine whether the stress associated with publicity surrounding that event might have lead to increased cancer rates in the years after 1978.

It will be important for public health officials to devise expanded research strategies for past and future environmental episodes that consider every plausible hypothesis for disease etiology. Latency periods prior to clinically apparent cancer may be very long. The promotion of already initiated cancer cells, through a stress-induced neuroendocrine mechanism, is not a far-fetched idea. Parenthetically, we should remember that only three or four decades ago many doubted that tobacco smoke could cause lung cancer and that the environmental pollution during an intense London fog in 1952 could have been responsible for some 4,000 deaths. As we consider the plausibility of the stress explanation in the report by Hatch et al., it is important to observe that a temporary promotion of cancer should lead to a short period of increased risk, followed by a compensatory period of decreased risk. The decreased risk occurs because new cancers are not initiated by a promotional event. Promotion can only hasten the progression of an existing cancer to the stage when the disease becomes clinically apparent. Perhaps mortality rates for cancer will need to be added to such studies.

The theories of initiation and progression of cancer are complex, but in general, a promotional event will not produce new cancers, although it may appear to do so in the short run. The data presented by the authors are consistent with such a pattern. As the authors clearly recognize, it is also difficult to rule out the possibility that increased clinical surveillance, perhaps stimulated by the same anxiety that produced the stress, could produce a short-term increase followed by a compensatory decrease in cancer cases diagnosed among the study population. The fact that the initial increase did not occur until three years after the nuclear accident lends more support to the promotion hypothesis than the increase-in-clinical-surveillance hypothesis as the explanation for this phenomenon. Whatever the final explanation for this cancer cluster following the Three Mile Island nuclear accident, the cluster underscores the need for more epidemiologically driven studies of major and minor environmental accidents. The investigators who conduct such studies should keep an open mind to the potential etiological mechanisms by which the circumstances around us may cause human diseases, including cancer. ☐

Dwight T. Janerich, DDS, MPH

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Address reprint requests to: Dwight T. Janerich, DDS, MPH, Professor of Epidemiology, Department of Epidemiology and Public Health, Yale University School of Medicine, 60 College Street, New Haven, CT 06510.

Erratum

In: Kahn HS, Williamson DF, Stevens JA: Race and weight change in US women: the roles of socioeconomic and marital status. Am J Public Health 1991;81:319–323.

The abstract was improperly copy edited by staff, resulting in the deletion of some important information in the Results section. The Results section of the abstract should read as follows:

"After multiple adjustments, Black race, education below college level, and becoming married during the follow-up interval were each independently associated with an increased mean weight change. Using multivariate logistic analyses, Black race was not independently associated with an increased risk of major weight gain (change $\geq +13$ kg), but it was associated with a reduced likelihood of major weight loss (change ≤ -7 kg) (odds ratio = 0.64 [95% CI = 0.41, 0.97]). Very low family income was independently associated with the likelihood of both major weight gain (OR = 1.71 [95% CI = 1.15, 2.55]) and major weight loss (OR = 1.86 [95% CI = 1.18, 2.95])."

The Journal regrets the error.

Asthmatic Responses to Airborne Acid Aerosols

Bart D. Ostro, PhD, Michael J. Lipsett, MD, JD, Matthew B. Wiener, PharmD, and John C. Selner, MD

ABSTRACT

Background: **Controlled exposure studies suggest that asthmatics may be more sensitive to the respiratory effects of acidic aerosols than individuals without asthma. This study investigates whether acidic aerosols and other air pollutants are associated with respiratory symptoms in free-living asthmatics.**

Methods: **Daily concentrations of hydrogen ion (H^+), nitric acid, fine particulates, sulfates and nitrates were obtained during an intensive air monitoring effort in Denver, Colorado, in the winter of 1987–88. A panel of 207 asthmatics recorded respiratory symptoms, frequency of medication use, and related information in daily diaries. We used a multiple regression time-series model to analyze which air pollutants, if any, were associated with health outcomes reported by study participants.**

Results: **Airborne H^+ was found to be significantly associated with several indicators of asthma status, including moderate or severe cough and shortness of breath. Cough was also associated with fine particulates, and shortness of breath with sulfates. Incorporating the participants' time spent outside and exercise intensity into the daily measure of exposure strengthened the association between these pollutants and asthmatic symptoms. Nitric acid and nitrates were not significantly associated with any respiratory symptom analyzed.**

Conclusions: **In this population of asthmatics, several outdoor air pollutants, particularly airborne acidity, were associated with daily respiratory symptoms. (*Am J Public Health.* 1991;81:694–702)**

Introduction

Exposure to acidic aerosols and sulfates has been associated with increased mortality and respiratory morbidity in a variety of settings.[1–5] Controlled exposure chamber studies suggest that individuals with asthma are more sensitive to the bronchoconstrictive effects of acidic sulfates than individuals without known pulmonary disease.[6–9] While the effects of exposure to ambient particulates (which may include acid aerosols) on asthmatic symptoms have been studied epidemiologically, previously published investigations have not included concurrent daily measurements of ambient airborne acidity.[4,5,10–12] Since the US Environmental Protection Agency has recently begun to consider whether to promulgate a national ambient air quality standard (NAAQS) for acid aerosols, it is important to assess the effects of relatively low levels of exposure on individuals in typical, free-living conditions. Thus, the need to extrapolate from studies of animals or of volunteers in controlled human clinical studies is obviated. In this study we report the results of an analysis of the association between ambient airborne acidity and daily respiratory symptoms recorded by a panel of asthmatics in Denver, Colorado during the winter of 1987–88.

Particulate air pollution has been associated with increased asthmatic symptoms and respiratory hospital admissions.[4,5,10–12] In an analysis of daily diaries of 454 asthmatics in the Los Angeles area, Whittemore and Korn[10] reported increased probabilities of attacks on days with high oxidant and total suspended particulate (TSP) pollution. Interpretation of this study is limited, however, by (among other things) the inadequacy of TSP as a measure of respirable particulate exposure and by significant covariation of TSP and ozone. Using a modification of the analytic model developed by Whittemore and Korn,[10,11] Perry, *et al*,[12] investigated the effects of fine (<2.5 microns in aerodynamic diameter) and coarse (>2.5 and <15 microns) particulates on respiratory symptoms, peak flows, and bronchodilator usage in 24 asthmatics in Denver in winter 1979. Fine nitrates were associated with increased symptoms and bronchodilator usage, while fine sulfates were associated with increased bronchodilator usage. The implications of this analysis are unclear because of the small sample size, the relatively low levels of particulate pollution, and the absence of acid aerosol measurements. In an ongoing study in Ontario, Canada, summer hospital admissions for asthma have been associated with exposure to ozone, sulfates, and sulfur dioxide.[4,5] These constituents of the "acid summer haze" were highly correlated; however, recent evidence[5] suggests that ozone may be less important than sulfates since there was a low correlation between respiratory admissions and ozone during one month of particularly high ozone concentrations. Again, acid aerosols were not measured.

The respiratory effects of airborne particulates on asthmatics and others have been attributed to the irritant, acidic frac-

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Address reprint requests to Bart D. Ostro, PhD, California Department of Health Services, 2151 Berkeley Way, Annex 11, Berkeley, CA 94704. Dr. Lipsett is also with the Department of Health Services at Berkeley; Dr. Wiener is with Pharmatech Inc., Englewood, CO. Dr. Selner is with the Allergy Respiratory Institute of Colorado, Denver, CO. This paper, submitted to the Journal October 24, 1989, was revised and accepted for publication August 13, 1990.

TABLE 1—Characteristics of Patient Population Participating in the Denver Asthma Study

Characteristics	Total Study Population (n = 256)	Diary Participants (n = 207)
Mean age	44.3	45.6
% Male	34	34
Race (% White/Black/other)	94.6/2.3/3.1	96.5/1.5/2
Education (%)		
College	62	62
High school	37	37
Less than high school	1	1
Employment status (% employed outside home/homemaker/retired/other)	71/10/15/4	70/10.5/13.5/6
Mean subjective asthma severity rating (0 = none, 4 = incapacitating)	1.78	1.74
% Daily Theophylline	56.4	58.0
% Daily oral steroids	21.8	18.9
% Current smokers	3.1	3.0

tion of this heterogeneous mixture of pollutants.[13-15] In controlled chamber studies of healthy individuals and asthmatics, acidic sulfates (H_2SO_4 and NH_4HSO_4) have been associated with bronchoconstriction and increased bronchial reactivity.[6-9] In a study involving short exposures (30 minutes at rest followed by 10 minutes of exercise), Koenig, *et al*,[6] reported significant changes in flow rates and pulmonary resistance in adolescent asthmatics exposed to 100 $\mu g/m^3$ H_2SO_4. Some studies showed effects only at higher levels of exposure which exceed typical urban peak sulfate levels,[7-9] and some showed no effects.[16] Nitrates and nitric acid have been studied less than acidic sulfates. Overall, the results of chamber studies suggest that asthmatics may be an appropriate "sensitive population" to study epidemiologically in relation to airborne acidity.

An intensive daily air monitoring program in the metropolitan Denver area was conducted from November 1987 to February 1988 to investigate the chemical composition and particle size distribution of the winter "brown cloud." The Integrated Environmental Management Program of Region VIII of the US Environmental Protection Agency (EPA) conducted a complementary air monitoring program from November 1987 to March 1988. The joint monitoring effort included twice-daily measurements of particulates (sulfates, nitrates, and all particles less than 2.5 microns in diameter or PM2.5), airborne acidity, and sulfur dioxide at two sites. Aerosol acidity was measured using annular denuder instrumentation[17] and included nitric acid vapor and nanoequivalents of hydrogen ion (H^+) in particulate form, which was not further speciated (into H_2SO_4 or NH_4HSO_4). Meteorologic data were collected by the State of Colorado and by Stapleton Airport in Denver. This period of intense air monitoring provided an occasion to investigate whether exposure to fine acidic particulates is associated with respiratory effects in asthmatic individuals. No previously published epidemiologic studies have correlated concurrently measured airborne acidity with asthmatic symptoms and behavior.

Data and Methods

Study Population

Study participants were recruited from patients attending the clinic of one co-author (JCS). Diagnosis of asthma was made in each case by history and signs of airway obstruction on physical examination, confirmed by spirometric demonstration of obstruction reversible with a β-agonist bronchodilator (>15 percent change in FEV_1). Asthmatic patients were identified by clinic staff and were recruited for participation either during an office visit or by telephone and postcard contact. Denver residents between ages 18 and 70 were eligible to participate in the study if they had asthma currently managed with medication. Individuals with any other chronic medical condition that would restrict their activity were excluded.

This panel of asthmatics was asked to record, on a daily basis for several months, information about symptoms, medication use, utilization of medical services, indoor exposures, and other variables described below. In the context of the air monitoring efforts noted previously, a panel study has several advantages over other investigative designs. First, it provides a large number of observations, increasing the degrees of freedom and the stability of any estimates. Second, problems of confounding, omitted variables, and exposure assessment, while always present, are substantially reduced since individuals serve as their own controls over time. Thus, the impact of factors that vary daily, such as air pollution, can be isolated while other factors are held constant. Third, the consideration of a given sample in one metropolitan location over time eliminates the potential for any intercity confounding.

Health Measurements and Covariates

After giving informed consent, participants were required to fill out an intake questionnaire providing background data on demographics (age, sex, race, level of education, employment status, residential history), asthma severity and characteristic triggers and symptoms, medical history including medication use, smoking history, and previous environmental exposures. A diary instrument was designed to provide daily information on asthma symptoms, including the presence and severity on a scale of 0 to 4 (0 = none, 1 = mild, 2 = moderate, 3 = severe, 4 = incapacitating) of cough, wheeze, shortness of breath, chest tightness, and sputum production, as well as physician and emergency room visits. In addition, information was obtained on frequency of medication use, time spent outdoors, levels of exercise intensity and whether exercise occurred indoors or outside, indoor exposure to respiratory irritants (such as gas stoves, fireplaces, and environmental tobacco smoke), and occupational exposures. Participants were told that this was an investigation of environmental factors affecting asthma, but not that the principal variables of interest were air pollutants.

There were 330 intake questionnaires distributed (93 during office visits and 237 by mail) from November 15 to December 16, 1987. Of the initial group recruited, 256 returned the intake questionnaires. Study subjects were contacted by clinic staff intermittently throughout the study period (December 1987 through February 1988) to enhance compliance and continued participation. Two hundred and seven patients returned their daily diaries. There were similar distributions of demographic variables and asthma severity in the initial study population and those who returned the daily diaries (Table 1). The group un-

TABLE 2—Descriptive Statistics of Pollutant, Meteorological, and Health Variables

	N	Mean	Std Dev	Minimum	Maximum
Sulfates ($\mu g/m^3$)	74	2.11	2.47	0.12	12.65
Nitrates ($\mu g/m^3$)	54	6.83	17.56	0.01	24.5
PM2.5 ($\mu g/m^3$)	47	21.97	20.40	0.51	73.08
Sulfur dioxide ($\mu g/m^3$)	78	14.14	12.82	0.80	59.8
Nitric Acid ($\mu g/m^3$)	64	1.81	2.25	0.06	13.54
H^+ (neq/m^3)	27	10.07	9.52	2.00	41.31
SH+ (neq/m^3)	74	8.15	8.29	0.59	44.25
Minimum temp. (F)	105	17.71	11.97	−11.00	39.00
Asthma	105	16.33	3.84	7.41	30.56
Cough	105	15.08	4.26	7.41	32.00

SH^+ = Hydrogen ion readings including substitutions for missing values.

Asthma = Group probability (×100) of reporting a moderate or worse overall asthma condition on each day.

Cough = Group probability (×100) of reporting a moderate or worse cough on each day.

All pollutants are measured as the average from 9:00 am to 4:00 pm

TABLE 3—Correlation Coefficients for Pollution and Meteorological Variables

	SULF	SO_2	NITR	PM2.5	HNO_3	H^+	SH^+	Mintemp
SO_2	.51							
NITR	.20	.16						
PM2.5	.88	.64	.40					
HNO_3	.22	.37	.78	.68				
H^+	.66	.35	.40	NA	−.11			
SH^+	.86	.44	.15	.56	.23	1.0		
MINTEMP	−.45	.26	−.08	−.34	−.07	−.57	−.43	
RH	.35	.12	.15	.40	.17	.32	.34	−.34

SULF = Sulfates

SO_2 = Sulfur dioxide

NITR = Nitrates

PM2.5 = Particulate matter less than 2.5 microns in diameter

HNO_3 = Nitric acid

H^+ = Hydrogen ion

SH^+ = Hydrogen ion with substitution for missing values (see text)

Mintemp = Minimum daily temperature

RH = Relative humidity

NA = Not available

der study was predominantly White, female, employed, and well educated, with an average age of approximately 46. The mean subjective asthma rating of moderate to severe is supported by the relatively large proportion of individuals taking daily oral theophylline and steroid preparations.

Exposure Measurement

The ambient air pollutants incorporated in the analysis were daily measures of sulfates, nitrates, PM2.5, nitric acid, hydrogen ion (H^+), and sulfur dioxide. Two different pollutant averaging times were considered initially: the 24-hour average (ending at 4:00 pm on the reporting day) of pollutant readings from the monitor in downtown Denver, located about two miles from the clinic; and the daytime average (9:00 am to 4:00 pm) from the same monitor. Since these averaging times were highly correlated (usually $r > .90$), only the daytime average was used in subsequent analysis. A second monitor, located about seven miles away in the northwestern suburb of Arvada, also provided measurements of these pollutants. Table 2 provides descriptive statistics for the pollutant, meteorologic and health variables. We found no statistically significant relationship between any pollutant and either the number of the day of the survey period or whether the reading was on a weekday or weekend and no significant time trend for either daily minimum temperature or daily humidity.

There was substantial covariation among some of the pollutants; the highest correlations were observed between PM2.5 and sulfates ($r = .88$), and between nitrates and nitric acid ($r = .78$). Sulfates and H^+ were both negatively correlated with maximum daily temperature (Table 3). Of the pollutants measured on a daily basis, sulfates had the most complete data. Values for missing days of PM2.5 and nitrates were derived from stepwise regressions using all other pollutants. PM2.5 could be predicted well by concurrent values of sulfates and nitrates ($r = .97$), and nitrates could be predicted by nighttime (4:00 pm to 9:00 am the following day) PM2.5 readings ($r = .86$).

The data on H^+ required additional attention. Because of problems in sample processing, the H^+ data were available only for half of the study period. The first seven days of H^+ readings were extremely high, with a mean of 288 nanoequivalents per cubic meter (neq/m^3), versus a mean of 10.1 for the following 28 days. The hydrogen ion concentrations during the first seven days were also unrelated to any other pollutant or meteorologic variables, while during the next 28 days there was a correlation, as expected, between H^+ and sulfates ($r = .66$). Consequently, the first seven days were dropped from the subsequent analysis. For H^+, missing values for the entire survey period were predicted using regression results with concurrent sulfate as the explanatory variable. With these substitutions, the daytime mean of H^+ was 8.15 neq/m^3 or approximately 0.4 $\mu g/m^3$ measured as sulfuric acid, with the highest daily average of 44 neq/m^3. All subsequent analysis for H^+ combined actual and imputed values. The one-hour peak may be three to four times higher than this longer-term average. These levels are typical of urban areas.[18] Analysis of the two monitors recording H^+ indicated that the concentrations of airborne acidity were fairly evenly distributed throughout the Denver area; the daily correlation of the readings was 0.88. Thus, the one downtown monitor was used to represent daily levels of H^+. Ozone concentrations, which frequently are highly correlated with sulfates and acids in the summer, were essentially at background levels through the winter in Denver (i.e. the maximum one-hour concentration was 0.042 ppm) and do not correlate with H^+ concentrations.

Previous research[19,20] suggests that incorporating information on exercise and time spent outdoors will improve the estimate of exposure. Therefore, we developed a measure of exposure that involved the product of the ambient pollution level, the duration of exposure, the time spent indoors and outdoors, and the ventilation rates based on exercise level (i.e. ventilation rates of 7, 10, 25, and 40 liters per minute for resting, mild, moderate, or strenuous exercise, respectively). Since the actual time and duration of exercise was not recorded in the diary, it was assumed to occur during the day and to last for one hour. Hours outdoors and exercise levels were taken directly from the daily diary, while the penetration ratio of indoor to outdoor levels of H^+ of 0.3 was derived from the published literature.[21] This adjustment for exposure was crude but represents an improvement over simply using outdoor levels of air pollution. The extent of the reduction in measurement error (and improvement in exposure assessment) is specifically tested in the subsequent analysis.

All data from questionnaires, daily diaries, air quality, and meteorologic monitoring were coded and edited in a SAS format for analysis.

TABLE 4—Regression Coefficients Relating Ambient Air Pollution and Exposure-adjusted Air Pollution to Moderate Asthma Symptoms, Corrected for Autocorrelation (standard error in parentheses)

Independent Variables	Cough	Asthma Rating	Cough	Asthma Rating
Constant	5.9	0.026	−0.15	−7.9
Symptom on previous day	74.8^{a}	83.0^{a}	76.6^{a}	97.1^{a}
	(14.5)	(8.1)	(15.6)	(8.3)
Log (H^+)	0.76^{b}	0.31		
	(0.38)	(0.42)	—	—
Exposure-adjusted log (H^+)	—	—	1.09^{b}	1.00^{c}
			(0.51)	(0.57)
Day of Survey	−0.067^{c}	0.005	−0.76^{b}	0.014
	(0.038)	(0.035)	(0.041)	(0.036)
Weekend	−0.85^{c}	−0.04	−1.5^{b}	−0.39
	(0.66)	(0.79)	(0.71)	(0.82)
Gas stove	6.53^{a}	6.97^{a}	8.25^{a}	6.69^{a}
	(1.75)	(1.88)	(1.94)	(2.03)
Minimum temperature	0.012	0.003	0.055	0.029
	(0.039)	(0.044)	(0.044)	(0.050)
R^2	.44	.37	.46	.37
# of observations	7417	7420	6364	6365
(# of people)	(168)	(168)	(163)	(165)

NOTE: a) significant at $p < .01$; b) significant at $p < .05$; c) significant at $p < .10$. All coefficients are ×100.

Statistical Methods

For this analysis, two different health outcomes were used as dependent variables in the multiple regressions: the probability of a respondent reporting on a given day a moderate (or worse) cough and the probability of reporting a moderate (or worse) rating of overall asthma status. The daily diary question regarding asthma status follows queries about the severity of wheezing, shortness of breath, chest tightness, and sputum production and should, therefore, reflect the respondents' subjective overall ratings of their individual asthma symptoms. As a daily average, 16.33 percent of the respondents (range 7.4 to 30.6 percent) reported a moderate or worse asthma day and 15.1 percent (range of 7.4 to 32 percent) reported a moderate or worse cough (see Table 2).

We estimated effects on each health outcome using ordinary least squares regression models. Variables included in each regression were minimum daily temperature, and binary variables indicating daily use of a gas stove and if the response occurred on a weekend. Since reporting of symptoms decreased during the course of the study, we also included the time since the start of the study. The probability of a symptom on the previous day was included as an explanatory variable in light of previous reports.[10,11] Daily relative humidity was examined and eliminated from the model after it was consistently found to be unrelated to any of the health outcomes, as were several time-invariant variables such as sex and education. Each pollutant's effect on health was estimated by entering it separately into the regression equation.

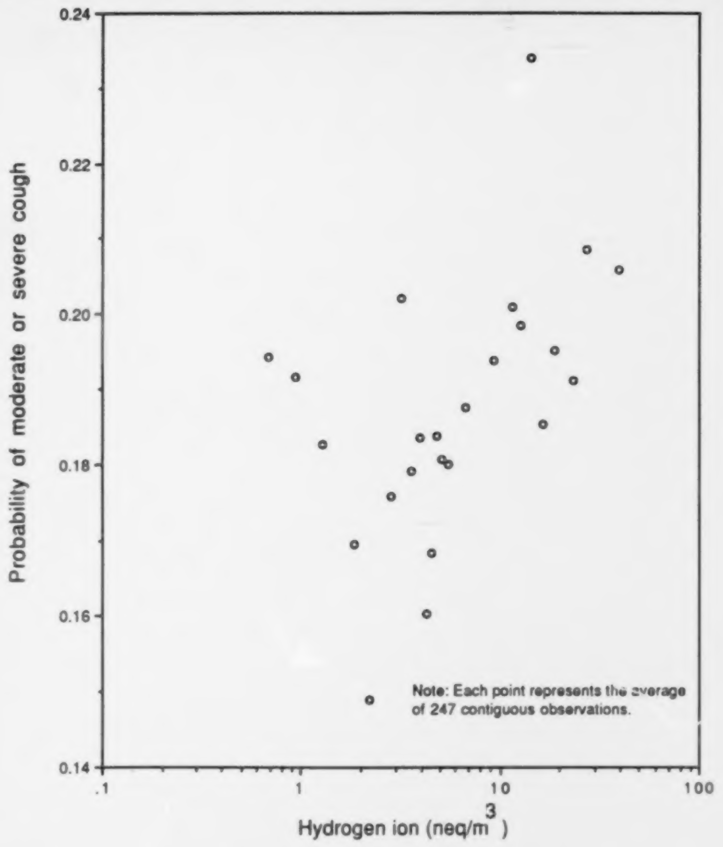
We also fitted a linear model of outcome using the logarithm of the pollutant, and a logistic model. There was little difference in model fit across the functional forms, but the log value of pollution provided the greatest association with the health outcomes. Thus, ordinary least squares with a log pollution term was used throughout the analysis. The similarity of results between ordinary least squares and logistic models probably was due to the relatively high rates of symptom incidence.

Significant autocorrelation was present, leading to biased estimates of the standard error of the estimated coefficients, and incorrect tests of significance.[22] The standard correction for autocorrelation involves estimation of a parameter (ρ) relating the regression residual at time t, r_t, to the residual for the previous time interval, r_{t-1}. When this feature is added to the estimation, the new model will have adjusted residuals which are uncorrelated. However, in this case, the procedure is biased and inconsistent because a lagged dependent variable (probability of symptom of the previous day) is included as an explanatory variable, and is correlated with the error term. Thus, the two-step procedure suggested by Fuller[23] was used. First, a predictive model was developed for the lagged value of the health effect (y_{t-1}) by regressing it on several independent variables, including time since start of the survey, specific air pollutants, minimum daily temperature, and humidity. Then the predicted value of y_{t-1} was used as an instrumental variable along with the other explanatory variables in the regression equation for y_t. Since the predicted value of the lagged dependent variable is uncorrelated with the adjusted residuals, the estimates are consistent. The estimation was performed using Proc Autoreg in SAS[24] with a maximum likelihood procedure that estimates both the regression coefficient and the autoregressive term.

Additional sensitivity analysis was conducted on the data, including a test for the influence of extreme observations, examination of other respiratory symptoms, and utilization of the exposure-adjusted measure of air pollution described above.

Results

Table 4 provides the ordinary least squares regression results for H^+, with the correction for autocorrelation. In the first two models, the dependent variables are the probability (×100) of reporting moderate or severe cough, and the probability of a moderate or worse asthma rating, respectively. The results indicate a statistically significant relationship between H^+ concentration and cough ($p < .01$) but not with overall asthma rating. Approximately 40 percent of the variation in the



Note: Each point represents the average of 247 contiguous observations.

FIGURE 1—Association of Moderate or Severe Cough with Hydrogen Ion

probability of reporting these symptoms was explained by the regression. For both cough and asthma status, reporting that symptom on the previous day was significantly related ($p < .0001$) with reporting a moderate or worse symptom. For cough, the day of the survey, responding on a weekend, and using a gas stove were statistically significant. Using overall asthma rating as the dependent variable, the only other significant risk factor (besides symptom on previous day) was usage of a gas stove ($p < .001$). For both cough and asthma status, autocorrelation corrections were appropriate and a second-order autocorrelation model was indicated. Figure 1 displays a plot of the relationship between hydrogen ion and moderate or severe cough.

Next, models with exposure-adjusted measure of H^+ were examined (Table 4). These models adjust outdoor air pollution levels for daily exercise levels and time spent outdoors. The results indicate that the exposure-adjusted measure of H^+ had a statistically significant association with both moderate cough ($p < .05$) and general asthma rating ($p < .10$). Figures 2 and 3 display these relationships.

Additional models correcting for autocorrelation were examined for each of the other pollutants, including sufates, sulfur dioxide, PM2.5, PM2.5 with substitutions for missing values, nitrates, and nitric acid. Each pollutant was considered individually and regressed first on cough and then on asthma rating and on shortness of breath. The other explanatory variables used in the multiple regression were those used above. Thus, Table 5 gives the results of these 18 separate regressions, each with one pollutant included in the equation. For comparison, the previous results for H^+ and exposure-adjusted H^+ are also included in the table.

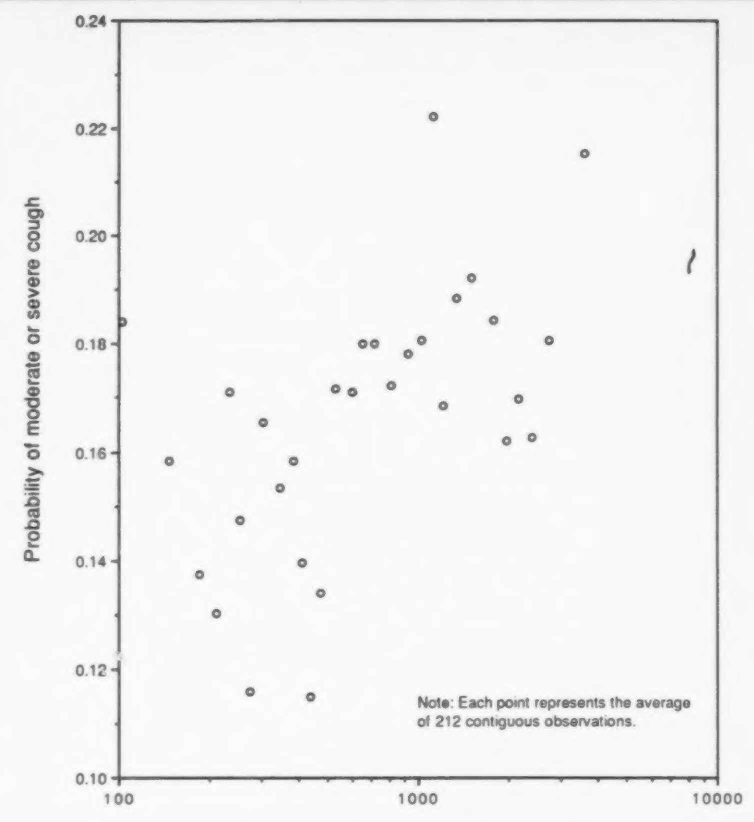
The results suggest that H^+, besides being associated with cough and overall subjective asthma rating, is also related to the reporting of moderate or worse shortness of breath. Besides H^+, the only other associations between individual air pollutants and these symptoms include: 1) sulfates and the probability of reporting moderate or worse shortness of breath ($p < .05$); and 2) PM2.5 and asthma rating ($p < .10$), after substitutions for missing values of PM2.5 were made, increasing the study duration from 46 to 75 days. Sulfur dioxide, nitrates, and nitric acid were not associated with moderate cough, asthma, or shortness of breath. The results for pollutants other than H^+ did not change when exposure-adjusted measures were used.

Additional analyses of the relationship of H^+ and asthmatic symptoms were conducted. The sensitivity of the models to the influence of extreme observations was tested using statistics proposed by Belsley, *et al.*[25] The test indicated that only the data from day = 2 had a relatively large influence on the regression results. The autoregressive model was rerun with that observation deleted, resulting in estimated regression coefficients that were essentially unchanged. Next, one- and two-day lags in H^+ were examined, indicating that contemporaneous measures of pollution provided the best association with asthma status. Finally, several time-invariant factors including education, sex, and whether an individual was recruited through office visit or by telephone were individually examined in the regression model. Their inclusion did not alter the estimated pollution effects.

Discussion

This analysis tested for an association between respiratory symptoms and several different air pollutants in a group of generally moderate to severe adult asthmatics. The results indicate a statistically significant association between ambient concentrations of hydrogen ion and the proportion of the subjects reporting moderate or worse cough and shortness of breath, with a lesser association observed in relation to the study subjects' overall asthma rating. Several lines of statistical and biomedical evidence implicate hydrogen ion as the principal pollutant of concern. First, although hydrogen ion was moderately correlated with minimum temperature and humidity (see Table 3), neither of these potentially confounding meteorologic variables was associated with the respiratory effects under investigation. Thus, it is unlikely that the estimated effects of hydrogen ion should be attributed to meteorologic changes over the survey period.

Second, confounding due to exposure to ambient ozone, which may cause



Note: Each point represents the average of 212 contiguous observations.

FIGURE 2—Association of Moderate or Severe Cough with Exposure-Adjusted Hydrogen Ion

exacerbations of asthma, was not a factor here. In studies linking health effects in asthmatics with resulting from summer haze in Southern Ontario[4] and particulate matter in Los Angeles,[10] ozone was highly correlated with the particle measure used in each study (sulfates and total suspended particulates, respectively), which in turn could be a surrogate for acid aerosol. In the Denver winter of 1987–88, ozone was generally found at background levels, as expected, with a maximum hourly concentration of 0.042 ppm.

Third, sensitivity analysis indicates that the findings were robust to functional form, specification, and influential observations. Linear and logistic functional forms provided similar results, and the estimated effect of air pollution was not affected by the exclusion of influential observations. In addition, the regressions were corrected for autocorrelation. Therefore, the results of the above analysis indicate that the symptoms reported are not likely to be due to a statistical anomaly.

The estimated impact of air pollution is greater when an adjustment for actual exposure is undertaken. For example, as indicated in Table 4, after adjusting exposure for intensity and location of exercise, time spent outdoors, and penetration rate of H^+, the estimated effect of air pollution on cough is 43 percent greater than that obtained from the unadjusted pollution measurements. This suggests that dose-response estimates for H^+ that do not incorporate behavioral factors affecting actual exposures may seriously underestimate the impact of air pollution.

Evidence from controlled exposure studies supports an etiologic role of hydrogen ion exposure in evoking asthmatic symptoms. Increasing sulfate aerosol acidity enhances the bronchoconstrictive effects in asthmatics in rough proportion to the acidity.[7] Exposure to 75 $\mu g/m^3$ of the nonacidic respiratory irritant, ferric sulfate, did not cause adverse effects on pulmonary mechanics or respiratory symptoms in normal or asthmatic subjects.[26] More recently, Fine, *et al*,[27] showed that buffered acid aerosols were significantly more potent bronchoconstrictors than unbuffered acids at the same pH, suggesting that the total available hydrogen ion is a critical determinant of airway responsiveness. The relationship between asthmatic symptoms and pulmonary function changes due to acid exposure is not clear-cut; however, some studies indicate that symptoms (especially cough) precede accompanying changes in respiratory mechanics, although this has not been consistently observed.[27–29]

This study suggests that hydrogen ion is the pollutant of primary concern. Although both sulfates and hydrogen ion were significantly related to the probability of reporting moderate or worse shortness of breath, only exposure-adjusted hydrogen ion was associated with the probability of reporting both a moderate or worse cough and asthma rating. Concentrations of particulate matter less than 2.5 microns were mildly associated with overall asthma ratings, but not with either cough or shortness of breath. Supporting evidence comes from a study of both asthmatics and healthy individuals,[30] in which ambient sulfates were not associated with acute respiratory symptoms, but airborne pH levels were. The importance of hydrogen ion over total sulfates or particle mass has also been suggested by the US EPA.[31] Since sulfate and particle measurements may include both acidic and nonacidic exposures, they incorporate measurement error which would bias downward the estimated regression coefficients and level of significance.

Of additional interest is the lack of association between nitrates or nitric acid and respiratory symptoms. Limited data from controlled chamber studies suggest that brief (10–16 min) exposure to high concentrations (up to 1,000 $\mu g/m^3$) of inorganic nitrate ($NaNo_3$) aerosols do not evoke respiratory symptoms or changes in pulmonary function in mild asthmatics.[32,33] Thus, there is some (admittedly sparse) clinical evidence to support the notion that particulate nitrates *per se* would not be expected to cause the effects reported here.

Our findings ostensibly conflict with those of Perry, *et al*,[12] who reported a mild association between respiratory symptoms and nitrates, but not sulfates, in a sample of 24 asthmatics. There may be at least three reasons for this discrepancy. First, the small sample size limits the generalizability of their findings. Second, the

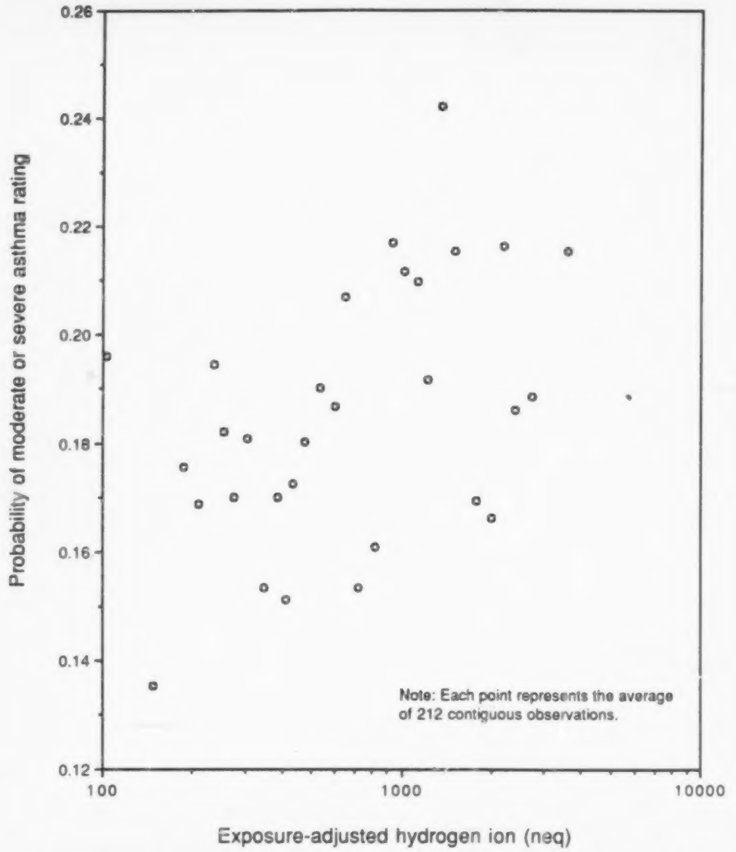


FIGURE 3—Association of Moderate or Severe Asthma Rating with Exposure-Adjusted Hydrogen Ion

rate of respiratory symptom reporting in that investigation was very low relative to ours. Third, there is a difference in pollutants under consideration and their dispersion patterns. Perry, *et al*, measured nitrates, sulfates, sulfur dioxide, and PM2.5 but not acid aerosols. In the Perry study and the present study the correlations among sulfates, sulfur dioxide, PM2.5, and temperature were remarkably similar, but nitrates were less correlated with other pollutants in the present study. This suggests that the sources of nitrates may have been quite different. Prior to the use of nylon back-up filters in the mid-1980s, nitrates were commonly measured inaccurately. Population growth and changes in automobile travel patterns also may have led to a change in exposures. In addition, during the winter of 1987–88, the major electric utilities, as part of a clean air campaign, burned natural gas instead of coal for alternate 15-day periods. These factors may have altered the dispersion pattern of nitrates, relative to sulfates and other pollutants, and the subsequent exposures of residents of the metropolitan area.

Since only about half of the individuals in this study completed their diaries every day, we investigated the potential for bias due to variable daily response rates. We found no relationship between air pollution and the number of people responding on a given day. We also investigated whether health status influenced response rates and found no relationship between overall severity of asthma and the number of days that people filled out the daily diaries. We infer that responses recorded in the daily diaries represent those of the group as a whole.

In this study population, the products of gas consumption appeared to exacerbate asthmatic symptoms. Peak indoor exposures to nitrogen dioxide (NO_2) resulting from gas stove use have been reported to range from 0.2 to 1.7 ppm,[34] which includes concentrations at which effects on symptoms or bronchial reactivity in asthmatics have been reported.[35,36] The concentrations of NO_2 and other constituents of natural gas combustion were not measured, but it is plausible that some symptoms were attributable to such indoor exposures.[37]

The results of this study, while supporting an association between asthmatic symptoms and ambient hydrogen ion exposure, should be interpreted with caution since pollutant measurements from fixed-site monitors cannot accurately represent individual exposures. Further, asthmatics may change their activity or medication use in order to reduce or eliminate symptoms related to air pollution; this would reduce the estimated effect of air pollution. Adjusting the pollution exposures for exercise and time outdoors increased the magnitude and precision of the estimate.

This study has several strengths, including the prospective nature of the investigation, the relatively large sample size, and the screening of the survey participants. Unlike previous large-scale studies of asthmatics, the sample for this study was recruited from one private practice at one location. Medical and personal histories were fully documented. Since the study was conducted as a time-series in one city over a relatively short period of time, geographic confounding factors were minimized, as was potential confounding due to ozone exposure. Relevant airborne acidity data, never previously collected concurrently with asthmatic symptom data, were also available.

This study provides evidence from a free-living population that is qualitatively consistent with the results of several controlled clinical investigations. Furthermore, our analysis indicates that, among moderate to severe asthmatics, respiratory symptoms occur at average hydrogen ion concentrations substantially lower than those at which effects on pulmonary mechanics have been observed in the laboratory.[7,9,28,29] That we were able to detect acute symptoms at the ambient concentrations measured was somewhat surprising. However, the study population was in general more severely afflicted than asthmatics volunteering for chamber studies. Some patients with severe asthma are known to have airway mucus with low pH, which may increase their susceptibility to acidic environmental insults.[38] Furthermore, the Denver winter temperatures and low relative humidity may have potentiated the effects of the airborne acidity (although we did not observe any obvious independent or interactive effects of these factors in the regression analysis). Cold dry air has been reported to enhance

TABLE 5—Estimated Regression Coefficients, Corrected for Autocorrelation, for Pollutants Examined Individually (standard error in parentheses)

Pollutants*	Asthma	Cough	Shortness of Breath
H^+	0.31	0.76^a	1.05^a
	(0.42)	(0.38)	(0.42)
Exposure adjusted H^+	1.00^b	1.09^a	1.33^a
	(0.57)	(0.51)	(0.57)
Sulfate	0.24	0.42	0.77^a
	(0.39)	(0.35)	(0.38)
SO_2	0.26	0.27	0.04
	(0.38)	(0.33)	(0.38)
PM2.5	0.06	0.12	−0.35
	(0.53)	(0.43)	(0.52)
PM2.5(Subs)	0.06^b	0.14	0.29
	(0.03)	(0.29)	(0.38)
Nitrate	−0.25	−0.21	0.10
	(0.25)	(0.94)	(0.25)
Nitric acid	−0.03	−0.32	−0.10
	(0.35)	(0.34)	(0.34)

a) significant at $p < .05$; b) = $p < .10$

PM2.5(Subs) = substitute for missing values of PM2.5 = 8.67 • sulfate + 1.071 • nitrate

*All pollutants are in log terms. Besides the air pollutant, all equations include day of the survey, symptom on previous day, an indicator for weekend vs. weekday, use of gas stove, and minimum temperature as explanatory variables and are corrected for autocorrelation.

All coefficients are ×100.

bronchoconstriction caused by sulfur dioxide in one of the few experiments investigating interaction between physical and chemical exposures.[39] In other contexts, observational studies have found air pollution effects at concentrations below those at which chamber studies reported positive results.[40] This suggests that factors that are typically not included in chamber studies may play an important role in determining individuals' responses to air pollution. ☐

Acknowledgments

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Homeless Mentally Ill Will Benefit from Research Demonstration Projects

A unique collaborative effort to improve housing and services for homeless people who have a severe mental illness was announced recently by Health and Human Services Secretary Louis W. Sullivan, MD. The Department of Housing and Urban Development will join with HHS in supporting six research demonstration projects, each of which will combine housing assistance with mental health services.

The new projects will be carried out in five cities across the country and will be administered by the National Institute of Mental Health.

According to previous NIMH-sponsored research, one third of the homeless population has severe and persistent mental illnesses, including schizophrenia and mood disorders. Yet with adequate housing, treatment and support, most of these individuals can lead productive lives in the community.

NIMH funding for the new projects will be approximately $5.8 million in FY 1991 out of a planned total of $16.8 million over the next three years. HUD will contribute over $10 million in housing assistance. Projects will be located in Boston, Massachusetts; San Diego, California; Baltimore, Maryland; Cincinnati, Ohio; and two in New York City. In *Boston,* the project will evaluate two alternatives to housing homeless mentally ill individuals; consumer-managed group homes and independent living arrangements will be compared to determine which approach results in better housing stability with higher levels of functioning.

In *San Diego,* the research demonstration project will evaluate the effectiveness of linking comprehensive services to independent housing for severely mentally ill homeless persons.

In *Baltimore,* the project will test an "assertive community treatment" program designed to increase the clients' access to housing, as well as general health and mental health services.

The project in *Cincinnati* will compare the benefits of transitional housing and permanent housing.

In *New York City,* one research demonstration project will help individuals on the street obtain housing, such as single-room occupancy units, made available through a city/state agreement, while another project will be directed at mentally ill men who are preparing to leave a homeless shelter.

For further information, contact the Department of Health and Human Services, (301) 436-4536.

A Randomized Trial to Evaluate the Risk of Gastrointestinal Disease due to Consumption of Drinking Water Meeting Current Microbiological Standards

Pierre Payment, PhD, Lesley Richardson, MSc, Jack Siemiatycki, PhD, Ron Dewar, MSc, Michael Edwardes, PhD, and Eduardo Franco, PhD

ABSTRACT

Background: This project directly and empirically measured the level of gastrointestinal (GI) illness related to the consumption of tapwater prepared from sewage-contaminated surface waters and meeting current water quality criteria.

Methods: A randomized intervention trial was carried out; 299 eligible households were supplied with domestic water filters (reverse-osmosis) that eliminate microbial and chemical contaminants from their water, and 307 households were left with their usual tapwater without a filter. The GI symptomatology was evaluated by means of a family health diary maintained prospectively by all study families over a 15-month period.

Results: The estimated annual incidence of GI illness was 0.76 among tapwater drinkers compared with 0.50 among filtered water drinkers ($p < 0.01$). These findings were consistently observed in all population subgroups.

Conclusion: It is estimated that 35% of the reported GI illnesses among the tapwater drinkers were water-related and preventable. Our results raise questions about the adequacy of current standards of drinking water quality to prevent waterborne endemic gastrointestinal illness. (*Am J Public Health.* 1991;81:703–706)

Introduction

It has long been considered that water disinfected at 0.5 mg/L of free residual chlorine for 30 minutes at a pH less than 8, with a turbidity of less than 1 NTU would constitute minimal health risk to consumers.[1] Recently, however, increasingly sensitive analytic methods have permitted the detection of viruses[2] and parasites[3] in water meeting current water quality standards. This and several reported outbreaks of viral and parasitic diseases of waterborne origin[4] have prompted a review of the safety of current standards.[5] The suggested standards are based on the absence of fecal coliform bacteria in drinking water, fecal coliform bacteria being indicators of fecal pollution, and the ensuing risk of the presence of pathogenic microorganisms. The absence of total coliform bacteria in a majority of the samples collected is also recommended as an indicator of treatment efficiency.

While the current microbiological standards of water quality have virtually eliminated explosive outbreaks of waterborne disease,[6] there is no assurance that populations consuming drinking water prepared from sewage-contaminated wastes are free of less virulent gastrointestinal (GI) disease. While studies have addressed the health effects of recycled wastewaters[7,8] or of substandard drinking water,[9] all have relied on cases reported to the health care system to evaluate potential health effects. Other studies have addressed the health risks associated with chemicals in water or with disinfection byproducts.[10]

The present project was intended to directly and empirically answer the following question: Is there any measurable excess of GI illness related to the consumption of tap water prepared from sewage-contaminated surface waters and which meets current microbiological and physico-chemical water quality criteria?

Methods

Study Area

The study was carried out in a suburban area of Montreal comprised mainly of French Canadians with socioeconomic and education levels similar to the average of the Metropolitan Montreal area. The residents of this area are served by a single water treatment plant, which uses state-of-the-art treatment with pre-disinfection, flocculation by alum, rapid sand filtration, ozonation, and final disinfection by chlorine or chlorine dioxide. The raw water to be treated is drawn from a river which is mainly contaminated by human sewage discharges with little contamination from chemicals. The distribution system serves a population of about 40,000 people and distributes water meeting current microbiological and physico-chemical water quality standards.[5] A pilot study had shown that over 90 percent of the population drinks unmodified tap water and that the water was considered by the consumers to be of good to excellent quality.

Study Design

In order to determine the level of gastrointestinal illnesses attributable to drink-

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Address reprint requests to Pierre Payment, PhD, Centre de recherche en virologie, Institute Armand-Frappier, Université du Québec, 531 Boul. des Prairies, Laval, Québec, Canada H7V 1B7. Coauthors are with the Centre de recherche en épidémiologie et médecine préventive, at the Université du Québec. This paper, submitted to the Journal October 5, 1990, was revised and accepted for publication February 19, 1991.

ing water, we carried out a randomized trial in which some households were left with their usual tap water and others were supplied with a domestic water filter which eliminates most chemicals and all particulate and microbiological matter from the tap water.

The sampling process was undertaken to enroll approximately 600 households for study, half randomly assigned to the filtered water group and half to the tap water group. Using a directory of inhabited addresses for the study area, households were selected at random and telephoned in random order. A screening questionnaire served to determine whether the household satisfied the five eligibility criteria: 1) owner-occupied; 2) French-speaking occupants; 3) regular consumers of tap water (as opposed to bottled or filtered water); 4) at least one child between the ages of 2 and 18 living in household; 5) willingness to participate in a longitudinal trial in which a random half of the households would have a filter installed. A total of 3,741 households were contacted before we achieved our quota of eligible study households. After randomization, there were 299 families in the filtered water group (1,206 individuals) and 308 families (1,201 individuals) in the tap water group.

The selection of eligible families was completed in January 1988 but the full-scale morbidity follow-up started in March 1988 after all filtration units had been installed in the appropriate households. Health outcomes of interest were monitored until June 1989. Because of the difficulty of monitoring during the summer vacation period there was a two-month gap in the surveillance during July and August 1988. We will refer to the March 1988 to June 1988 observation period as Period 1, and September 1988 to June 1989 as Period 2.

Intervention

For the tap water group participants, there was no intervention and no suggestion to modify their normal water consumption patterns.

For the participants allocated to the filtered water group, under-the-sink filtration units were installed by professional plumbers, at no expense to the family, during the start-up period of February-March 1988. The filters were commercial domestic reverse-osmosis units attached to the cold water line and placed under the kitchen sink. They were composed of a prefilter, an activated charcoal filter, a reverse osmosis membrane cartridge, and a 15-liter reservoir containing the purified water. During the reverse-osmosis process, tap-water is forced through a molecular filter which is permeable only to very small molecules such as water and some inorganic salts. Microorganisms, particulates, and most chemicals are rejected by the membrane and discharged to the drain. The filters were obtained, installed, and maintained by a local supplier. Prefilters and charcoal were changed after nine months of use or upon clogging. Efficiency of the units was monitored weekly by the family informant using a small conductivity tester. Water in the reservoir of such units is pressurized by the water pressure of the main inlet to the house and is available at all times in sufficient quantity for normal usage by the whole family.

Background Information and Health Outcomes

All participating households were initially visited in January and February 1988 by a nurse who carried out an interview on baseline factors and explained the way the health data would be collected. The baseline data included sociodemographic and medical history information on each household member. The health end-point of interest was GI symptomatology and this was determined primarily by means of a family health diary maintained prospectively by all study families for the whole period except for the two-month summer break. A simple self-administered questionnaire was designed to be completed in two-week blocks. One household member, usually the female head, was designated as the informant. The informant was responsible for noting for any household member on each calendar day each incident of the following events: nausea, vomiting, or diarrhea (soft or liquid). If any of these three symptoms was recorded, the presence of the following accompanying symptoms was to be noted: fever, cramps, muscular pain, cold or flu, sore throat, absence from work or school, visit to a doctor, or hospitalization. For any positive reports, we inquired as to possible reasons such as food overindulgence, pregnancy, and visits to countries with high endemic levels of GI illnesses. At the end of each two-week period, a nurse from our staff telephoned the informant and asked for the information recorded on the diary. The regular and frequent contact by our staff was intended to encourage reporting accuracy.

For the purpose of statistical analyses reported here, we defined an episode of highly credible GI (HCGI) symptoms as involving at least one of the following combinations: 1) vomiting or liquid diarrhea with or without confinement to bed, consultation with a doctor, or hospitalization, or 2) nausea or soft diarrhea combined with abdominal cramps with or without absence from school or work, confinement to bed, consultation with a doctor, or hospitalization. Episodes with plausible etiologies apart from the one under study (e.g. pregnancy) were excluded from consideration. An episode was further defined as consisting of one or more symptomatic days, with at least six consecutive symptom-free days between the episodes.

Statistical Analyses

Outcome parameters included the percentage of study subjects with episodes, the annual incidence of episodes and the duration of episodes. The study subjects can be seen as the family units or as the individuals within these family units.

Some analyses were crude and thus included no other covariates apart from group membership while others included covariates to adjust for some imbalance in the two groups. The covariates included in the model, in addition to tap water/filter water status were: age, sex, and geographic subregion. Estimates of incidence of HCGI illnesses in the two groups were derived by means of Poisson regression methods.[11] Since this technique counts each episode as an independent event, a correction had to be made for the correlation between repeat episodes by the same unit of observation. This was done using the "quasi-likelihood" approach,[12] which amounts to reducing the nominal significance of each result by a factor related to the amount of this correlation. Adjusted rates were estimated from the results of the fitted model. Each regression model gave an estimate of the p-value associated with the parameter for group membership. The preventable fraction of disease was computed by dividing the difference in rate between the tap water and filtered water by the rate in tap water group. The percentage of subjects with episodes over some period of time was estimated by two methods, logistic regression, and proportional hazards analysis. Both of these techniques are based on data relating to the first episode that the individual experienced during the period of observation; the number of episodes experienced is ignored. The logistic regression analysis[13] requires that the period of

TABLE 1—Distribution of Selected Individual Characteristics in the Two Study Groups at the Study Outset

	% in Group:	
Sub-groups	Filtered Water (N = 1206)	Tap Water (N = 1202)
Age group		
0–5	9.6	6.2
6–20	37.4	39.4
21–49	46.0	43.2
50+	7.0	11.2
Sex		
Female	49.4	49.5
Male	50.6	50.5

NOTE: There were some dropouts during the course of the study. The total sample sizes at the beginning of the second period were 1,194 in the filtered water group and 1,067 in the tap water group. The distributions of characteristics were almost identical in the second period to those shown above in the first period.

observation be equal for all study subjects. Thus this analysis had to be limited to subjects for whom there were no missing data during the entire field work period. However, this analysis provided estimates of percentages of study subjects in both study groups who had at least one episode. The proportional hazards analysis by contrast can use all study subjects irrespective of the length of observation period but it estimates a rate ratio rather than a pair of rates.[14]

Water Consumption

To evaluate whether subjects in the two study groups were in fact consuming the type of water we intended them to consume, we carried out a survey to estimate each person's water consumption habits during one week day and one weekend day in the preceding seven days. This questionnaire asked about water consumed inside and outside the home and listed different types of use (e.g. straight, in juice, in hot drinks). From this survey, we computed an estimate of weekly water consumption inside and outside the home for each person. The response rate to this supplemental self-administered questionnaire was 82 percent among tap water subjects and 84 percent among filter water subjects.

Microbiological Analysis of Water

Water from the filtration plant (raw and treated) from 25 sampling sites in the distribution system and filtration units was sampled at least weekly for bacteriological and physicochemical analysis. The bacteriologic parameters estimated were: heterotrophic plate counts at 20°C and 35°C, total coliforms and fecal coliforms; *Clostridium perfringens, Aeromonas hydrophila, Pseudomonas aeruginosa.* Large volume virological analysis of the raw and treated water from the plant were also performed at least weekly according to the methods described previously.[15] Physicochemical parameters evaluated were: pH, turbidity, and residual disinfectant concentration. All analyses were performed by the Montréal Laboratory of the Ministry of Environment of Québec according to *Standard Methods for the Examination of Water and Wastewater*[16] and to the Ministry recommendations.[17]

Results

Water Quality

River water used for the preparation of the drinking water was contaminated by fecal material of human origin as indicated by the high levels of the following pathogens and indicator microorganisms found in the raw water entering the plant: human enteric viruses (78 viruses/100L), total coliforms (57,530 cfu/L), fecal coliforms (3,674 cfu/L), *C. perfringens* (623 cfu/L), *A. hydrophila* (6,590 cfu/L), and *P. aeruginosa* (41 cfu/L). Treated water (tap water) at the filtration plant was free of any detectable indicator bacteria and free of human enteric viruses and in compliance with regulations. However, bacterial growth was observed in the water filtration units as previously reported.[18]

Study Population Characteristics

Table 1 reports the age and sex distributions of the two study groups at the study outset. By and large, the randomization produced very similar groups, although there were slightly more very young and fewer very old in the filtered water group than in the tap water group.

Over 95 percent of eligible families maintained their participation for Period 1 (Table 2). For the entire two periods, the participation rates were 86.0 percent in the tap water group and 92.6 percent in the filtered water group.

Water Consumption

Table 3 shows that at home the two groups consumed about the same amount of water, the tap water group consuming only tap water and the filtered water group consuming almost entirely filtered water. Away from home, both groups consumed mainly tap water and in both groups there was about five times as much water consumed at home than outside the home.

Disease Outcome

During the study period there were no outbreaks of gastrointestinal diseases officially reported to the local health authorities. The number of visits to physicians for gastrointestinal symptoms and of hospitalizations was similar in both groups.

Altogether 65 percent of families experienced at least one episode of HCGI symptoms during the entire period of observation and about 23 percent of individuals had at least one episode. The overall incidence of HCGI was 0.66 episodes per person per year and the mean duration of an episode was 1.9 days. The yearly incidence of HCGI symptoms was higher in Period 1 (Table 4A) than in Period 2 (Table 4B). Age was a strong determinant of incidence of HCGI symptoms with the highest rates in the under 5 age group.

Table 5 reports the incidence rates in the two groups adjusted to the covariate distribution of the tap water group. Three units of observation were analyzed: the family as a whole, the informant, and the youngest child within the family. Again, the pattern of difference between tap and filtered water groups was quite similar for the different units of observation and across the two periods. Using the entire family data, the preventable fraction was 35 percent in the Period 1 and 32.8 percent in the Period 2. Combining the two periods, the adjusted incidence rates were 0.76 in the tap water group and 0.50 in the filtered water group ($p < 0.01$).

The HCGI episode is a synthetic combination of a number of GI-related symptoms. Similar analyses were done for each of the symptoms included in the diary: the findings were very similar to what was observed for the synthetic HCGI variable.

Dose-Response Relationship

To determine whether increased tap water consumption in the home was associated with increased risk of morbidity, we restricted attention to the tap water group. Poisson regression analyses were carried out with the number of episodes in the entire second period as the dependent variable and the estimated number of glasses of water consumed per week as an independent variable. Table 6 shows that for both informants and youngest chil-

TABLE 2—Family Participation Patterns in the Two Study Groups across the Two Study Periods

Outcome	Filtered Water Number of Families	Tap Water Number of Families
Period 1—March to June 1988		
Total number enrolled	299	308
Quit—problems with filter	6	—
Quit—changed to filtered or bottled water	—	9
Quit—moved out of study area	5	3
Quit—unwilling to continue	3	1
Completed Period 1	285	295
Summer break		
Changed to ineligible[a]	2	9
Period 2—Sept 1988 to June 1989		
Total number enrolled	283	286
Quit—problems with filter	2	—
Quit—changed to filtered or bottled water	—	10
Quit—moved out of study area	3	5
Quit—unwilling to continue	1	6
Completed Period 2	277	265

[a] Most of these were due to a youngest child leaving home or reaching 18 years of age.

TABLE 3—Mean Number of Glasses of Water Consumed Weekly by Type of Water and by Group[a]

Type of Water Consumed	Group Filtered Water	 Tap Water
At home		
Filtered water	37.4	0.0
Tap water	1.9	37.8
Away from home		
Filtered water	0.3	0.0
Tap water	7.7	8.7

[a] Based on a separate survey of the study population. Responses were obtained from 957 subjects in the filtered water group and 891 in the tap water group.

dren, there was an increasing risk with increasing consumption. Among the tap water group there was no relationship between the quantity of water consumed outside the home and morbidity. Among the filtered water group there was no relationship between the quantity of water consumed inside or outside the home and morbidity (results not shown).

Nature of the Excess

Analyses were also carried out to determine whether the higher incidence in the tap water group was due to more people experiencing symptoms or whether it was due to the same core of people experiencing more episodes. These analyses were restricted to those families for whom complete data were available for the entire study. As shown in Table 7, there was a slight difference between the two groups in the percentage of subjects with episodes, but the number of episodes per person affected was considerably higher in the tap water group than in the filtered water group. This implies that the main effect of tap water consumption was to increase the number of episodes among susceptibles.

Discussion

This randomized trial was carried out to determine whether drinking water conforming to currently accepted microbiological standards was the source of any excess of GI disease when compared with a group drinking the same water after a molecular filtration to remove contaminants. The study was carried out in a relatively homogeneous middle class suburban neighborhood, and the overall observed incidence of gastrointestinal symptoms was similar to incidence rates reported in other North American populations.[19–21] The difference between the periods is within range for this type of symptoms. There was, however, a clear and significant excess of HCGI symptoms among the tap water drinkers as compared with filtered water drinkers. This finding was bolstered by the relative consistency of the effect across different age, sex, subregion, and informant/proxy categories, and the observation that among tap water drinkers there appears to be an increasing risk with increasing consumption.

Before proceeding to speculate as to the reasons and impact of these findings, it is important to consider possible sources of bias, the main one being reporting bias. Although the intervention was randomized, study subjects were not blinded. The possibility of having a sham filter installed in homes of tap water subjects was considered, but it was not technically and financially feasible. Subjects with a sham filter would have perceived the inefficacy of their filter and would have recognized the characteristic chlorine odor of tap water.

It is conceivable that subjects participating in the tap water group but aware of the study objectives exaggerated their morbidity relative to the filtered water group. Similarly, subjects in the filtered water group, if they had faith in the quality of their water may have underestimated their morbidity. However, we doubt whether such biases had any significant effect. Symptoms such as vomiting and liquid diarrhea are not subjective and open to interpretation. If such biases were operating, it is likely that they would be most pronounced at the outset and diminish with time as the novelty of the health diary wore off. However, over the 15-month study period, the relative excess of GI disease was fairly constant, arguing against any such bias. Furthermore, the observation of a dose-response relationship is less plausibly attributable to such biases. As a general check on the credibility of the quality of reporting, it is of interest that the estimated incidence rate is similar to reports from other North American centers and the relative consistency of the effect across the different categories.[19–21]

Outside the home, study subjects from both groups consumed water from all kinds of sources. Thus there was some misclassification in labeling one group as tap water consumers and the other as filtered water consumers. Moreover, this would tend to attenuate the effect of any true difference due to domestic water source and thus reinforces the importance of the difference observed. Similarly, the finding of an apparent dose-response relationship in the tap water group is particularly impressive given the crudeness of the measure of dose which was based on a one-point-in-time survey. The treated water leaving the filtration point was free of bacterial pathogens and indicator bacteria as well as of cultivable human enteric viruses.

TABLE 4A—Incidence of HCGI Episodes for Period 1 for the Two Study Groups[a]

Subgroups	Filtered Water N	Filtered Water Rate[b]	Filtered Water S.E.[c]	Tap Water N	Tap Water Rate	Tap Water S.E.
Age group						
0–5	116	1.66	0.28	74	2.48	0.44
6–20	454	0.77	0.10	475	0.95	0.11
21–49	558	0.54	0.09	524	0.97	0.13
50+	85	0.31	0.25	135	0.46	0.24
Sex						
Male	613	0.75	0.09	609	0.88	0.10
Female	600	0.68	0.09	599	1.11	0.12
Informant/proxy						
Informant	298	0.65	0.13	304	1.00	0.16
Proxy	915	0.74	0.09	904	0.99	0.10
All subjects	1213	0.72	0.08	1206	0.99	0.10

[a] HCGI: highly credible gastrointestinal symptoms. See Methods Section for operational definition of HCGI and of an episode.

[b] Annual incidence per person.

[c] Standard error corrected for cluster nature of sampling.

TABLE 4B—Incidence of HCGI Episodes for Period 2 by Selected Population Subgroups[a]

Subgroups	Filtered Water N	Filtered Water Rate[b]	Filtered Water S.E.[c]	Tap Water N	Tap Water Rate	Tap Water S.E.
Age group						
0–5	115	1.13	0.18	72	1.59	0.27
6–20	428	0.46	0.07	427	0.66	0.08
21–49	512	0.39	0.06	458	0.56	0.08
50+	78	0.23	0.12	126	0.33	0.11
Sex						
Male	572	0.44	0.06	545	0.62	0.07
Female	561	0.52	0.07	538	0.68	0.08
Informant/proxy						
Informant	281	0.45	0.08	275	0.62	0.10
Proxy	852	0.49	0.06	808	0.65	0.07
All subjects	1133	0.48	0.06	1083	0.65	0.07

[a] HCGI: highly credible gastrointestinal symptoms. See Methods Section for operational definition of HCGI and of an episode.

[b] Annual incidence per person.

[c] Standard error corrected for cluster nature of sampling.

TABLE 5—Incidence of HCGI Episodes Derived from Poisson Regression Analysis, by Study Group for Each of Three Units of Observation[a]

Unit of Observation	Period 1 Filtered (n = 298)	Period 1 Tap (n = 304)	Period 1 Preventable Fraction (%)	Period 2 Filtered (n = 282)	Period 2 Tap (n = 276)	Period 2 Preventable Fraction (%)
Family	0.65	1.00**	35.0	0.43	0.64**	32.8
Informant	0.64	1.00*	36.0	0.44	0.63*	30.2
Youngest child	0.86	1.24*	30.7	0.56	0.80*	30.0

[a] Annual incidence per unit of observation, derived by Poisson regression with age, sex, and subregion in the models.

*Difference between the incidence rates is significant at $p < 0.05$

**Difference between the incidence rates is significant at $p < 0.01$

Public Health Impact

In the past decade, there has been growing concern about the safety of public water supplies. Much of this concern has focused on presumed noninfectious outcomes such as cancer and birth defects and, in this context, there is some fear that chlorination byproducts may be harmful.[10,22] Heretofore, it was assumed that current water quality standards were sufficient to protect the public against the risk of GI disease; the apparent lack of outbreaks seemed to support this view. Our results suggest that there is a non-trivial endemic level of unreported GI diseases due to the consumption of tap water; 35 percent of gastrointestinal illness was estimated to be attributable to the consumption of drinking water meeting current water quality standards.[5]

The main impact of tap water consumption seemed to be in increasing the numbers of episodes of susceptibles rather than in increasing the numbers of people with episodes. The severity of these illnesses was low and resulted in few demands on the health care system. The community studied is typical of many North American communities in terms of its water source being contaminated by sewage discharges, its modern water treatment facilities, its treatment procedures which are widely used, and its sociodemographic characteristics. Our results raise questions about the adequacy of current standards of drinking water quality, including source water quality, treatment procedures, and bacteriological quality of the water in the distribution system as well as in-house plumbing. ☐

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TABLE 6—Incidence of HCGI Episodes in Period 2 by Amount of Tap Water Consumed among Tap Water Subjects[a]

Unit of Observation/ Consumption	Number		Incidence[b]
Informant			
1–25 glasses/week	24		.42
26–45 glasses/week	92		.70
46+ glasses/week	114		.66
Trend test		p = .10	
Youngest child			
1–25 glasses/week	74		.55
26–45 glasses/week	92		.89
46+ glasses/week	64		1.10
Trend test		p = .10	

[a] Annual incidence per unit of observation during Period 2.
[b] Poisson regression model including age, sex, sub-region.

TABLE 7—Percent of Study Subjects Reporting HCGI Symptoms and Mean Number of Episodes per Unit of Observation in Both Periods Combined

	Group			
	Filtered Water (n = 272)		Tap Water (n = 262)	
Unit of Observation	% with Episodes[a]	Mean Number of Episodes[b]	% with Episodes	Mean Number of Episodes
Family	62.0	3.82	67.7	4.81
Informant	20.0	1.70	23.1	2.10
Youngest child	42.3	1.83	46.3	2.37

[a] Derived by logistic regression with covariables age, sex, geographic sub-region. Only subjects who had complete data sets in both periods are included in this table.
[b] Mean number of episodes among those subjects who reported at least one episode.

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Risk Assessment and Control of Waterborne Giardiasis

Joan B. Rose, PhD, Charles N. Haas, PhD, and Stig Regli

ABSTRACT

Background: Waterborne giardiasis has been increasing in the United States with 95 outbreaks reported over the last 25 years. The Safe Drinking Water Act has mandated control of this pathogen.

Methods: A risk assessment model was developed to estimate risk of infection after exposure to treated waters containing varying levels of *Giardia* cysts. The model was defined by a dose-response curve developed from human feeding studies for *Giardia* and assumed 2L of water consumption per day. Data on concentrations and distribution of the organism in source waters were used to assess exposure after varying reductions achieved through treatment.

Results: In surveys reporting prevalence and levels of *Giardia* cyst contamination, average levels of cysts in surface waters ranged from 0.33 to 104/100L; from pristine watersheds (protected from all human activity) 0.6 to 5/100L. Yearly risks were 4.8×10^{-3} for systems using polluted waters and 1.3×10^{-4} for pristine waters with a 10^{-3} treatment reduction.

Conclusion: Public Health officials will need to work with the water industry to ensure a risk of less than 1/10,000 for source waters with 0.7 to 70 cysts per 100 liters through treatment achieving reduction of 10^{-3} to 10^{-5}, respectively, of *Giardia* cysts. (*Am J Public Health.* 1991;81:709–713)

Introduction

Giardia is the most frequently isolated enteric protozoan from populations worldwide and the most common pathogenic parasite in the United States.[1] Waterborne giardiasis has been increasing in the United States with 95 outbreaks reported over the last 25 years,[2] and *Giardia* is the most common identifiable etiological agent of all waterborne outbreaks. Bennet *et al.*[3] have also estimated that 60 percent of all *Giardia* infections are acquired through contaminated water.

Giardia cysts may be found in water as a result of the deposition of fecal material from both man and animals. Surveys of *Giardia* cyst levels in various waters indicate that 26–43 percent of the surface waters were contaminated with *Giardia* cysts ranging in concentrations from 0.3 to 100 cysts per 100L.[5-10] Sykora[9] has reported an average of 10^4 cysts/100L in raw sewage with an approximate reduction of 10^{-2} after treatment (152 cysts/100 L).

The Surface Water Treatment Rule has been promulgated to address the amendments to the Safe Drinking Water Act for controlling *Giardia* in treated drinking water.[4] This rule mandates that all surface waters be treated to achieve at least a reduction of 10^{-3} (99.9 percent removal) of *Giardia* cysts. Disinfection is required for all systems and filtration is required unless the system meets site specific criteria and has a protected watershed. The US Environmental Protection Agency (EPA) has also recommended that a treatment be provided to ensure that populations are not subject to risk of infection of greater than 1:10,000 (10^{-4}) for a yearly exposure, and that this is an acceptable level of safety for potable waters.

As the new regulations and rules come into effect, health departments will be called upon to ensure that the public is adequately protected against waterborne disease. Risk assessment is a tool by which health officials can communicate with the water industry by interpreting water quality surveys and assisting in defining the adequacy of treatment adhering to EPA's recommendations of potable water quality and acceptable public health risks. This will become particularly important as states implement the Surface Water Treatment Rule, evaluate new technologies, and determine what water management practices will impact public health.

This paper presents a risk assessment model that was used to estimate risk of infection due to waterborne exposure to *Giardia*. Dose-response curves were developed based on human infectivity studies, and data on the occurrence of *Giardia* cyst contamination in waters throughout the US were used to define the water treatment needed to reduce the risk of waterborne giardiasis.

Methods

To predict a potential public impact when the risk is small, models are used to estimate the risk after exposure. Haas reviewed three probability models for their ability to describe experimental dose-response data for humans after exposure to

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Address reprint requests to Joan B. Rose, PhD, Assistant Professor, Department of Environmental and Occupational Health, University of South Florida, 13301 Bruce B. Downs Blvd, Tampa, FL 33612-3899. Dr. Haas is with the Department of Environmental Engineering, Illinois Institute of Technology; Mr. Regli is with the Office of Drinking Water, US Environmental Protection Agency, Washington, DC. This paper, submitted to the Journal April 19, 1990, was revised and accepted for publication December 10, 1990.

various enteric microorganisms.[11] Using the same approach, the simple exponential and beta distributed effectiveness models were evaluated for prediction of *Giardia* infectivity using experimental data developed by Rendtorff[12,13] in human feeding studies.

In Rendtorff's experiments, *Giardia* cyst doses ranging from 1 to 10^6 were fed to volunteers, and a positive response was measured by cyst excretion in the feces (Appendix). Laboratory dose-response studies generally appear to be conducted under conditions where the distribution of microorganisms in the administered dose may be regarded as Poisson. Under these conditions, if one microorganism is sufficient to cause an infection, and if host-microorganism interactions are constant, then the probability of an infection (P_i) resulting from ingestion of a single volume of liquid containing an average number of organisms (N) may be defined by a simple exponential equation.

$$P_i = 1 - \exp(-rN)$$

In this equation, r is the fraction of microorganisms that are ingested which survive to initiate infections ("host-microorganism interaction probability") (Appendix). In this particular case, the exponential model was statistically consistent with the Rendtorff data and r was calculated to equal −0.01982.

The 95 percent confidence limits to the parameter r in the exponential model were computed using a likelihood ratio technique. The resulting interval estimate for r is 0.009798–0.03582. This range was used in the preparation of Figure 1.

Exposure Estimates

Surface waters were classified into two categories: polluted waters contaminated by sewage and agricultural discharges; and pristine waters originating from protected watersheds without point source pollution or input from human activities. *Giardia* cyst levels were examined for the peak level of contamination from a single sample, average concentrations for each site, and average cyst levels for each water classification. Geometric means were calculated from the average concentrations from each site.[5–10] *Giardia* cyst levels were calculated as cysts/100 L as large volumes of water are routinely sampled for determining levels of contamination.

Using the exponential model, the potential risk of infection was determined with varying levels of *Giardia* cysts in drinking water. The model assumed the consumption of two liters of water per day, and exposure N in the formula was defined by numbers of cysts per liter times two liters. Levels of *Giardia* cysts found in polluted and pristine source waters and assumed levels of 10^{-3}, 10^{-4}, and 10^{-5} removal by treatment were used to estimate exposure in the model. Maximum daily risk was estimated using the peak level of contamination and yearly risk was determined using the geometric mean concentration of *Giardia* cysts for 365 days of exposure. Probability of infection was determined assuming a Poisson distribution of microorganisms in the drinking water,[11] i.e.,

> risk of contracting one or more infections = $1-(1-\Pr(N))^x$ where x = the number of days exposed and Pr(N) = the daily risk using the geometric average for N.

Data from five waterborne outbreaks of giardiasis were also evaluated for the *Giardia* cyst level of contamination detected in the drinking water and the attack rates in the exposed population.[14] This information was compared to the estimated infection rates developed by the exponential model after varying days of exposure.

Results

To ensure less than a daily risk of 10^{-4}, systems using source waters containing 250, 2,500 and 25,000 cysts per 100L would need to reduce the level of *Giardia* cysts by 10^{-3}, 10^{-4} and 10^{-5}, respectively, through drinking water treatment (Table 1). Examples of yearly risks for exposure to varying levels of *Giardia* cysts in drinking water are summarized in Table 2. To ensure less than a yearly risk of 10^{-4}, systems using source waters containing a geometric concentration of 0.7, 7.0 and 70 cysts per 100L would require a reduction of 10^{-3}, 10^{-4} and 10^{-5}, respectively, by treatment.

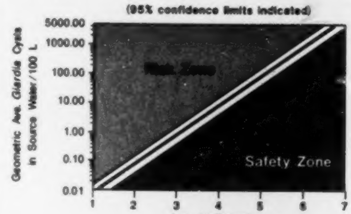

FIGURE 1—Log 10 Treatment Reductions for *Giardia* Cysts in Source Waters Yielding 10^{-4} Annual Risk (95% confidence limits indicated)

In surveys reporting prevalence and levels of *Giardia* cyst contamination in surface waters, average levels of cysts ranged from 0.33 to 104/100L. Eight of the areas where samples were collected from received treated sewage or agricultural discharges, however, the level of pollution was not documented.[5–10] In water samples originating from pristine watersheds (protected from all human activity) *Giardia* cyst levels averaged 0.6 to 5/100L. The percentage of positive samples did not vary dramatically from polluted to pristine waters (43 percent and 35 percent, respectively). Peak contamination in a single sample was 5.5 times higher from polluted waters than from pristine waters. However, geometric averages of cysts were 50 times higher for samples collected from polluted waters versus those collected from pristine waters.

The majority of data for the pristine waters was developed from a study by

TABLE 1—Estimated Daily Risk of *Giardia* Infections from Various Levels of Cyst Contamination in Drinking Water Using an Exponential Risk Assessment Model

Worst Daily Risk Pr(N)	Cyst Exposure	Maximum Cyst Concentration in Finished Water per 100L	Maximum Cyst Concentration in Source Water Per Level of Treatment Reductions		
			10^{-3}	10^{-4}	10^{-5}
				Cysts/100L	
$10^{-3.5}$	0.015	0.75[a]	7.5×10^2	7.5×10^3	7.5×10^4
10^{-4}	0.005	0.25	2.5×10^2	2.5×10^3	2.5×10^4
$10^{-4.5}$	0.0015	0.075	75	7.5×10^2	7.5×10^3
10^{-5}	0.0005	0.025	25	2.5×10^2	2.5×10^3

[a] Levels of cysts detected during waterborne outbreaks of giardiasis.

TABLE 2—Estimated Yearly Risk of *Giardia* Infections from Various Levels of Cyst Contamination in Drinking Water Using an Exponential Risk Assessment Model

Yearly Risk 1-(1-Pr[a] (N))^{365}	Average Daily Cyst Exposure for a Year	Geometric Mean Cyst Concentration in Finished Water per 100L	Geometric Mean Cyst Concentration in Source Water Per Level of Treatment Reductions 10^{-3} Cysts/100 L	 10^{-4}	 10^{-5}
$10^{-3.5}$	4×10^{-5}	2×10^{-3}	2.0	20	200
$10^{-4.0}$	1.4×10^{-5}	7×10^{-4}	0.7	7.0	70
$10^{-4.5}$	4×10^{-6}	2×10^{-4}	0.2	2.0	20
10^{-5}	1.4×10^{-6}	7×10^{-5}	0.07	0.7	7.0

[a] Pr = Daily Probability of Risk

TABLE 3—Potential Levels of Cysts in Treated Drinking Water and Probability of Infection

Source Water	Reduction by Treatment	Geometric Cyst Levels per 100L in Drinking Water	Yearly Risk[b]	With Peak Contamination Cysts/100L in Drinking Water	Daily Risk
Polluted	10^{-3}	0.033	4.8×10^{-3}	0.625	2.5×10^{-4}
33	10^{-4}	0.0033	4.8×10^{-4}	0.062	2.5×10^{-5}
cysts[a]/100L	10^{-5}	0.00033	4.8×10^{-5}	0.006	2.5×10^{-6}
Pristine	10^{-3}	0.0009	1.3×10^{-4}	0.114	4.6×10^{-5}
0.9	10^{-4}	0.00009	1.3×10^{-5}	0.0114	4.6×10^{-6}
cysts[a]/100L	10^{-5}	0.000009	1.3×10^{-6}	0.0011	4.6×10^{-7}

[a] geometric average
[b] Risk = 1−(1−Pr(N))^{365}

Ongerth, *et al.*[6] The Tolt, Green and Cedar watersheds in the State of Washington were extensively sampled with 222 samples collected. The averages reported in the publication included a 22 percent recovery adjustment for method efficiencies. The data from the Ongerth study[6] used in this paper did not include such adjustments since the data used from other studies[5,7–10] did not take into account adjustments for recovery efficiencies.

Daily and yearly risks were developed for the peak cyst level and the geometric average, respectively, for the two water categories (Polluted and Pristine). The data are shown in Table 3. Yearly risks ranged from 4.8×10^{-3} for systems using polluted waters with a 10^{-3} treatment reduction to 1.3×10^{-6} for pristine waters with a 10^{-5} treatment reduction. Between 10^{-4} and 10^{-5} reduction with treatment would be required for polluted waters to achieve a similar risk as pristine waters treated for a 10^{-3} removal of *Giardia* cysts. For daily risks for waters containing peak levels of *Giardia* cysts, the differences were not as dramatic between systems using polluted versus pristine source waters. The peak daily risk was approximately five times greater in systems using polluted versus pristine waters receiving similar treatment for removal of cysts.

Water samples were collected and *Giardia* cyst levels were determined during the investigation of five waterborne outbreaks of giardiasis. Attack rates varied from a low of 0.5 percent in the Houtzdale and Pittston outbreaks to a high of 16 percent in the Ft. Plain outbreak. The level of cyst contamination ranged from 0.6 to 21 cysts/100L. Generally the lower levels of contamination were associated with lower levels of infection in the population. The data are shown in Table 4.

All five outbreaks were associated with unfiltered chlorinated surface waters. Three factors primarily influenced the attack rate for infection: first, the level of contamination; second, the level of cyst viability and inactivation through chlorination; and third, the length of exposure to the population whose prior exposure to *Giardia* (and potential immunity) was unknown. Much of this detail is not known. The exponential risk model was applied to the known cyst levels from the outbreak and infection rates were estimated for varying days of exposure. It is unlikely that there was a single day of exposure or that the exposure was equal to the duration of the outbreaks. If one uses intermediate times of exposure, five to 10 days, the infection rates developed by the model ranged from 0.1 to 7.7 percent.

Discussion

Many regulatory and public health agencies have accepted risk assessment models to evaluate the importance of chemical pollutants in water. How well these models reflect reality depends on the accuracy of the model for characterizing the independent variables, assumptions, and the data used to develop the dose-response curves and exposure.

The most essential component to the risk model described in this paper is the dose-response curve. *Giardia* species/strains are known to have a low infectious dose. Rendtorff and Holt demonstrated in 1954 that ingestion of as few as 10 cysts was capable of initiating infection in two volunteers. It must be kept in mind that infection was measured by cyst excretion and illness was not determined. Asymptomatic *Giardia* infections may range between 39 percent and 76 percent for children less than five years of age and adults, respectively.[15,16] Symptomatic infections, however, have been reported at a rate of 50 percent to 67 percent and as high as 91 percent.[17] Chronic giardiasis may also develop in as many as 58 percent of the population infected.[18] Thus the illness to infection ratio is highly variable. We addressed only infection in this analysis of risk.

Another important issue regarding the dose response curve based on the Rendtorff data is uncertainty about infectivity due to strain variation and the immune response to infection by different populations. The Rendtorff data are derived from one *Giardia lamblia* strain and one relatively small sample population. The confidence interval around the probability of infection does not take these uncertainties into account when using the model as a predictive tool.

Assuming the dose response relationship derived from the Rendtorff data is representative, we may be overestimating

TABLE 4—*Giardia* Cyst Contamination of Drinking Waters during Waterborne Outbreaks of Giardiasis and the Relationship to an Exponential Risk Model

		Outbreak Data		Estimated Infection Rates[b] after Exposure for (Days)				
Outbreak	Duration (days)	*Giardia*[a] Cysts/100L	Attack Rates %	1	5	10	30	60
Pittsfield	30	21	3	0.84	3.9	7.7	21	38
Ft. Plain	60	19.5	16	0.78	3.8	7.5	21	37
Houtzdale	?	2.4	0.5	0.09	0.4	0.9	2.7	5.2
Pittston	?	0.87	0.5	0.03	0.1	0.3	0.9	1.8
Wilkes-Barre	?	11	9	0.4	2.0	3.9	11.3	21
		0.6	1	0.02	0.1	0.2	0.6	1.2

[a] Data received from Dr. Judith Sauch, Environmental Protection Agency, Cincinnati, OH.[14,15]
[b] Estimated from the exponential risk model.

risks based on the assumption that all cysts are viable and all cysts in water are species which infect humans. In addition, the 2L of tap water consumption may represent overestimation of exposure, depending on individuals consumption of other beverages. However, the underestimation of risk may be of greater concern due to underestimation of exposure by the inefficiencies of the methods for detection of *Giardia* cysts in water.

In spite of its limitations, the model can be used to examine waterborne outbreak data and disease surveillance data associated with various exposure routes. Epidemiological data, may be even more insensitive and inaccurate. Attack rates for waterborne outbreaks, defined by the number of illnesses in the exposed population, are developed using a variety of methods; they do not take into account unreported cases or asymptomatic infections, and may include individuals who were not exposed. Estimation of infection based on attack rates is probably understated.

Levels of cysts ranging from 0.6 to 21/100L detected in waters were associated with outbreaks of disease. *Giardia* cyst levels found in drinking waters averaged 0.19/100L during non-outbreak conditions and were rarely above 1/100L.[19] Based on the data in Table 4, it appears the model may be useful in estimating probability of infection. The greatest limitation of the model may be the underestimation of infection and disease through the use of cyst concentrations in water without taking into account method efficiencies, peak contamination levels, and duration of exposure.

By applying information on *Giardia* cyst levels to Figure 1, public health agencies, regulatory agencies, and water utilities can evaluate the adequacy of current treatment of a drinking water supply, the need for more effective treatment, or the type of treatment needed in developing new water supplies.

Public health workers must not only have an understanding of waterborne disease. The implications of new regulations, watershed protection, and water treatment on indigenous potential waterborne infections in the population must also be understood. Standard methods are available to assess *Giardia* cyst contamination in water. Health agencies can utilize such data in a risk assessment, cost-benefit approach to establish rational policies for public health protection. ☐

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APPENDIX

For the purpose of modeling, it was assumed that the cysts have a random distribution in water and that a Poisson distribution would govern the probability of any given exposure.[11] The risk of infection was defined as a function of the exposure (N organisms) and the interaction of the host and pathogen. The parameters characterizing this interaction (r) were determined using the human infectivity data set. The data set for the experimental infection and development of the model areshown in Table A-1. The single-hit exponential equation for the probability of infections is shown in Table A-2. The constant characterizing the host-cyst interaction with the best fit values to the data set is also presented. An average r value (that fraction of microorganisms that are ingested which survive to initiate infection) for the exponential model was computed by determining the value of r at each dose. The chi square goodness of fit test was used to evaluate the appropriateness of the model for the data set.The single hit exponential model was used toevaluate the risk of waterborne giardiasis. Low level contamination of drinking water with *Giardia* (0.5 cysts/100L)would result in a potential infection rate of 0.02 percent in the exposed population using the exponential model while at high levels of contamination, (100 cysts/100L) a 3.9 percent infection rate was determined (Table A-3).

TABLE A-1—Data Set for Experimental Infection for *Giardia* Cyst Exposure

Cyst Dose	Volunteers' Response: Positive[a]	Volunteers' Response: Negative	Percent with Positive Response	Predicted Number of Positives[b]
1	0	5	0	0.098
10	2	0	100	0.359
25	6	14	30	7.809
10^2	2	0	100	1.724
10^4	3	0	100	3
10^5	3	0	100	3
3×10^5	3	0	100	3
10^6	2	0	100	2

[a] Positive response denoted by cyst excretion.
[b] Based on function $p = 1-\exp(-0.0198N)$
Source: Rendtorff[12,13]

TABLE A-2—Best Fit Distribution Parameters for *Giardia* Infection Probability

Model	*Exponential Model*
Equation	$p = 1 - \exp(-rN)$
Constant(s)	$r = (-.01982)$
Expected probability of infection from an average dose of 1 cyst	0.0198
Number of infections in a population of 100, each exposed to 1 cyst	2

*95% confidence intervals for $r = 0.009798–0.03582$.
Source: Haas[11]

TABLE A-3—Potential Risk of Infection in *Giardia* Contaminated Drinking Water Using an Exponential Risk Assessment Model

Giardia concentration in water: Cysts/100L	*Giardia* concentration in water: Cysts/L*	Daily Risk	Percent Infection
0.5	0.005	2.0×10^{-4}	0.02
0.8	0.008	3.2×10^{-4}	0.032
1	0.01	4.0×10^{-4}	0.04
5	0.05	$2\ \times10^{-3}$	0.2
10	0.1	4.0×10^{-3}	0.4
15	0.15	6.0×10^{-3}	0.6
20	0.2	8.0×10^{-3}	0.8
50	0.5	2.0×10^{-2}	2.0
100	1	3.9×10^{-2}	3.9
200	2	7.7×10^{-2}	7.7

*Assume consumption of two liters of drinking water per day/

Lyme Disease: A Proposed Ecological Index to Assess Areas of Risk in the Northeastern United States

ABSTRACT

Background: Recent public awareness has resulted in a demand for information about ways to reduce the risk of acquiring Lyme disease.

Methods: Twenty-two school properties and recreational areas within a Lyme disease endemic area of central Monmouth County, New Jersey were evaluated for risk of transmission using an ecological index on the suitability, amount, and access to *Ixodes dammini* habitat by target human populations and the abundance of infected adult ticks.

Results: The characterization of tick habitat accurately predicted the elimination of 11 sites from concern. Of the remaining 11 sites, six were classified high risk and five as moderate risk. On-site tick surveys identified infected *I. dammini* adults at only four sites (three risk; one moderate risk).

Conclusions: These results indicate that the use of selected ecological parameters provides a cost-effective method to rapidly identify areas at risk for Lyme disease transmission. (*Am J Public Health.* 1991;81:714–718)

Terry L. Schulze, PhD, R. Chadwick Taylor, MBA, George C. Taylor, MPH, and Edward M. Bosler, PhD

Introduction

The extensive media attention recently afforded to Lyme disease has resulted in a dramatic increase in public awareness and demand for information regarding this tick-borne, spirochete-induced illness. The public is concerned about ways to reduce the risk of acquiring Lyme disease, including the relative safety of discrete geographical areas such as schools, recreational areas, and residential communities.

Early ecological studies in New Jersey identified areas of potential risk at the regional level through deer surveys[1] and serosurveys of canines.[2] As Lyme disease became more prevalent, areas at risk were determined by plotting case reports.[3] While local health officials could identify general areas at risk by randomly conducting tick surveys or retrospectively plotting case reports, a cost-effective method to prospectively identify specific geographical areas at risk for Lyme disease was unavailable. We propose here an index for use by public health and allied agencies which systematically assesses specific geographical areas for potential and actual risk of Lyme disease transmission through evaluation of key ecological parameters.

Methods

Study Sites

In response to increased public inquiries, the Freehold Township Health Department petitioned the New Jersey State Department of Health for assistance in evaluating the relative risk of Lyme disease transmission at all eight public school properties and 14 recreational areas within its jurisdiction. The public school properties ranged in size from 4.83 to 69.59 acres ($\bar{x}$ = 19.80 acres). The recreational areas ranged in size from 0.5 to 200 acres ($\bar{x}$ = 40.5 acres).

Identification of Tick Habitat

The northern deer tick, *Ixodes dammini* Spielman, Clifford, Piesman and Corwin is most prevalent in wooded areas, associated ecotones, and old fields in secondary woody succession.[4] Extensive surveys of potential tick habitats common to the Inner and Outer Coastal Plains of New Jersey where the majority of Lyme disease cases are reported[3] confirmed these associations (unpublished data). In upland areas of the coastal plain physiographic regions of New Jersey and other northeastern states, habitats generally yielding the highest populations of *I. dammini* are comprised of mixed hardwood, mixed hardwood/pine, and pine forests with a dense shrub layer dominated by highbush blueberry, *Vaccinium corymbosum* L.; lowbush blueberry, *Vaccinium angustifolium* Ait.; and huckleberries, *Gaylussacia* spp. with some mountain laurel, *Kalmia latifolia* L.; bayberry, *Myrica pensylvanica* Loisel; and catbrier or greenbrier, *Smilax* spp.[5]

Address reprint requests to Terry L. Schulze, PhD, Division of Epidemiology and Disease Control, New Jersey State Department of Health, CN 369, 3635 Quakerbridge Road, Trenton, NJ 08625. R. C. Taylor is with Freehold Township Health Department, Freehold, NJ; G. C. Taylor is with Preventive Medicine Services, Environmental Health Section, USA MEDDAC, Fort Monmouth, NJ; Dr. Bosler is with New York State Department of Health, Health Science Center, State University of New York, Stony Brook, NY. This paper, submitted to the Journal March 27, 1990, was revised and accepted for publication January 23, 1991.

TABLE 1—Summary of *I. dammini* Abundance and Infection Rates from Recreational Areas Yielding Ticks in Freehold Township, New Jersey

Surveyed Area	Number of Adult Ticks (60-Minute Survey)	Rate of Infection by *B. burgdorferi* (%)
A	48	41.7
B	88	57.1
C	28	50.0
D	24	33.3
NWS Earle	74–180	41.9

Tick Surveys

Tick surveillance would normally be scheduled only for those sites identified as having high or moderate risk of potential transmission based on the evaluation of habitat. To test the efficacy of the proposed index, all 22 sites were surveyed for the presence of ticks. Tick populations were evaluated during the period of peak adult activity in the fall by walking throughout suitable tick habitat at each of the subject sites for a 60-minute period.[4,6] At smaller sites, the survey time was reduced and the number of ticks collected was extrapolated to the 60-minute survey period for ease in comparison. This recognized survey technique provides an accurate representation of the expected number of adult ticks to be encountered by humans.[7] When no suitable habitat was present, the site was then randomly walked. Ticks adhering to investigators' light-colored, cotton coveralls were removed and subsequently evaluated for the presence of spirochetes using dark-field microscopy.[8] Based on previous studies in Monmouth County, the spirochetes identified in this study were assumed to be *Borrelia burgdorferi.*[9]

Ecological Index

The proposed ecological index is designed to use information on the type of plant communities and abundance of infected adult *I. dammini* to rapidly predict the risk of Lyme disease transmission at discrete geographical sites. Although plant communities may be assessed throughout the year, the seasonality of adult tick activity would naturally restrict the use of the proposed index in New Jersey to early spring and fall.[10] To alleviate this problem, the index was divided into two phases. Phase I permits the characterization of a site with respect to potential risk of transmission in the absence of tick activity by evaluating the suitability, amount, and accessibility of *I. dammini* habitat. Point values ranging from a minimum of 1 to a maximum of 5 are assigned to each of the three ecological parameters and totaled. The resulting score is used to select the appropriate response action for each site and determine the need to proceed to Phase II of the proposed index.

Rather than sole reliance on habitat to predict the risk of transmission at a particular site, Phase II of the proposed index is designed to utilize, in conjunction with the information on habitat, data on the presence and size of the deer tick population, and rate of *B. burgdorferi* infection. Point values of 1 to 5 are assigned to each of the five ecological parameters and totaled to achieve a final score for each site. This scoring procedure is used to categorize actual risk and establish priorities for intervention strategies.

Detailed descriptions of the ecological parameters, scoring procedures, and response actions for both phases of the proposed ecological index along with the survey form are provided in the Appendix in an instructional format.

Survey Procedure

The assessment of potential risk of transmission for each of the subject sites was performed on June 22, 1989 and July 19, 1989. To determine actual risk of transmission, the subject sites were surveyed for *I. dammini* between October 27, 1989 and November 8, 1989. These dates were well within the period of peak activity for *I. dammini* adults in New Jersey.[10] Surveys were performed only on fair days between 1000–1400 hours when peak diurnal activity was expected. Ticks collected during walking surveys were retained for subsequent analysis. Data on tick abundance and infection rate were recorded on the original survey form for each site. The sites were then ranked with respect to risk of transmission and need for intervention. To document that temporal and meteorological conditions were conducive to tick questing, a site at nearby Naval Weapons Station Earle (NWS Earle), Colts Neck, New Jersey exhibiting similar habitat and known to have consistently high populations of infected *I. dammini*[9] was also surveyed for ticks on each of the study dates.

Results

Potential Risk of Transmission

Of the eight public school properties, only one demonstrated high risk of transmission. Moderate risk of transmission was assigned to one school, while the remaining six schools received a low risk rating.

Regarding the recreational areas, five of 14 were considered to have high risk of transmission ratings, while four sites were assigned moderate risk of transmission ratings. The remaining five recreational areas received low risk of transmission ratings.

Actual Risk of Transmission

None of the eight school properties surveyed yielded *I. dammini* adults. On the survey days, the site at NWS Earle produced 88 and 114 ticks per 60-minute survey, confirming tick activity on the survey dates. Of the eight schools, one was classified as limited risk, three as no present risk, and four as no risk likely.

Ticks were collected from only four recreational areas. In 60-minute surveys, the four sites yielded between 24 and 88 adult ticks. At the NWS Earle site, 74, 88, and 180 ticks were collected in equivalent 60-minute surveys, again confirming tick activity. Minimum field infection rates at the four recreational sites ranged between 33.3 percent and 57.1 percent. During the survey period, 41.9 percent of adult ticks collected from NWS Earle were infected with spirochetes. The tick collection and infection rate data for sites yielding ticks are summarized in Table. 1. As a result of tick surveys, two of the 14 recreational sites were classified as definite risk, three as potential risk, one as limited risk, three as no present risk, and the remaining five as no risk likely.

Discussion

The proposed ecological index was designed for use in upland areas of the coastal plain physiographic regions of the northeastern United States where the majority of Lyme disease cases occur.[11] In this investigation, the proposed index provided a rapid, systematic, and cost-effective method of eliminating from public health concern high use public property by assessing the ability of existing plant communities to serve as habitat for *I. dammini.* Once significant potential risk was determined by surveys of vegetation, the actual risk of transmission and appropriate response action were determined based on the abundance of infected adult deer ticks. Where the response action may include chemical control, the proposed index will assist the user in identifying areas requiring treatment, thereby reducing unnecessary acaricide applications.

Although tested on only 22 sites comprising 725 acres within a hyperendemic area of New Jersey, we believe that, with minor modifications, the proposed index can be used in other ecosystems and for other tick vectors. For example, the "suitability of habitat" parameter can be modified to reflect differences in vegetation utilization by deer ticks in the ecosystem under investigation. Although we found the collection of adult ticks to be the most cost-effective method in characterizing the deer tick population, other survey techniques may be substituted to evaluate subadult *I. dammini* or other tick vectors. In such cases, criteria of the "tick abundance" and "tick infection rate" parameters will require modification. The proposed ecological index is designed to be a flexible survey instrument that provides public health officials and others with a means to rapidly assess the potential and actual risk of Lyme disease transmission at various sites within their jurisdictions. ☐

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APPENDIX
Lyme Disease Survey Form
Site ____________________
Investigator(s)
Potential Risk ____________________ **Date** ____________
Actual Risk ____________________ **Date** ____________

1. Suitability of Habitat: 1 2 3 4 5
Comments ____________________
____

2. Amount of Habitat: 1 2 3 4 5
Comments ____________________
____

3. Accessibility of Habitat: 1 2 3 4 5
Comments ____________________
____

4. Tick Abundance: 1 2 3 4 5
Comments ____________________
____

5. Tick Infection Rate: 1 2 3 4 5
Comments ____________________
____

Final Score:
Potential Risk of Transmission ____________
Response Action ____________________
____

Actual Risk of Transmission ____________/____________/____________ = ____________
Response Action ____________________
____

(continued)

APPENDIX continued

Instructions for Use of Lyme Disease Survey Form

Index Parameters

Phase I of the ecological index prospectively assesses discrete geographical sites for the potential risk of Lyme disease transmission by evaluating three ecological parameters. In the absence of tick data, point values are assigned to reflect the degree to which each parameter enhances the likelihood of transmission. In addition to these three parameters, Phase II employs tick abundance and tick infection rate to categorize these sites for the actual risk of Lyme disease transmission. Point values were similarly assigned to these parameters to complete the assessment.

Parameter Number	Parameter Description	Point Value	
		High	Low
1	Suitability of Habitat	5	1
2	Amount of Habitat	5	1
3	Accessibility of Habitat	5	1
4	Tick Abundance	5	1
5	Tick Infection Rate	5	1

Description of Parameter

1. Suitability of Habitat

This parameter addresses the suitability of the habitat (plant community) at the subject site for supporting populations of *Ixodes dammini* ticks.

Point Value	Habitat Description
5	= mixed hardwoods or mixed hardwoods/pine forest with substantial (>75% coverage) shrub layer dominated by highbush blueberries, lowbush blueberries, and huckleberries with some laurels, bayberries, catbrier, saplings/seedlings of the aforementioned tree species, and associated ecotones (Type 1 habitat).
4	= same as above except with a modest (25%–50% coverage) shrub layer (Type 2 habitat).
3	= same as above except with a sparse (<25% coverage) shrub layer or fields in secondary woody succession (Type 3 habitat).
2	= meadows, overgrown fields, or forested wetlands.
1	= agricultural fields, athletic fields, lawns, etc.

2. Amount of Habitat

This parameter represents the amount of habitat (Types 1, 2, and/or 3) at the subject site able to support populations of *Ixodes dammini* ticks.

Point Value	Amount of Suitable Tick Habitat
5	= 80% to 100% of subject site
4	= 60% to 79% of subject site
3	= 40% to 59% of subject site
2	= 20% to 39% of subject site
1	= 0% to 19% of subject site

3. Accessibility of Habitat

This parameter provides an estimate of how accessible the suitable tick habitat at the subject site is to the human population at risk.

Point Value	Description of Accessibility
5	= suitable tick habitat is easily accessed or people are encouraged to access the subject site (picnic areas, hiking or exercise trails, campsites, place of residence, playgrounds, school yards, etc).
4	= suitable tick habitat adjacent to subject site and readily accessible.
3	= suitable tick habitat present, but access discouraged because of natural or man-made barriers (catbrier, multiflora rose, or other impenetrable vegetation; water; fences; etc.); the area is totally undeveloped; or the amount of suitable habitat is limited.
2	= access to suitable tick habitat restricted (military facilities, private clubs, etc).
1	= no access or no suitable tick habitat present.

(continued)

APPENDIX continued

4. Tick Abundance

This parameter addresses the relative number of *I. dammini* adults in habitats within the subject site. The number of *I. dammini* adults collected during 60-minute walking surveys (or extrapolated equivalent) will be used for this phase of the ecological index.

Point Value	Tick Abundance
5	= >30 *I. dammini* adults per survey
4	= 11–29 *I. dammini* adults per survey
3	= 6–10 *I. dammini* adults per survey
2	= 1–5 *I. dammini* adults per survey
1	= 0 *I. dammini* adults per survey

5. Tick Infection Rate

This parameter takes into account the percentage of *I. dammini* adults collected at the subject site that are infected with *Borrelia burgdorferi.*

Point Value	Tick Infection Rate
5	= >40% of *I. dammini* adults infected
4	= 30% to 39% of *I. dammini* adults infected
3	= 20% to 29% of *I. dammini* adults infected
2	= 10% to 19% of *I. dammini* adults infected
1	= 0% to 9% of *I. dammini* adults infected

Scoring and Response Actions—Potential Risk of Transmission

After point values are assigned to each of the three ecological parameters, the point values are totaled to achieve a final score for the subject site. The final score is then compared to the three point ranges presented below to determine the appropriate response action for the subject site.

The response action designations are to be used by the investigator to assign priorities to sites within his/her jurisdiction with respect to potential risk of Lyme disease transmission in the absence of any data on tick abundance and rates of infection by *B. burgdorferi.* This phase of the ecological index may also serve as a guide/provide justification for use of limited resources.

Scoring Range	Description of Risk	Response Action
11–15	High Risk	Place under routine tick surveillance and establish rates of infection.
6–10	Moderate [illegible]	Consider periodic tick surveillance and establishment of infection rates.
<6	Low Risk	No action necessary at this time or any time in the future.

Scoring and Response Actions—Actual Risk of Transmission

Point values relating to the habitat parameters (Parameters 1–3) are totaled first. This figure is recorded and separated from the point value for Tick Abundance (Parameter 4) by a slash (/). The point value for Tick Infection Rate (Parameter 5) is similarly separated from the remainder of the index by a slash. As such, a site receiving the maximum rating for each parameter would be recorded as 15/5/5. By recording the point values in this fashion, the ecological index becomes more descriptive, allowing the investigator to rapidly determine the relative importance of habitat, tick abundance, and tick infection rate in the assessment of the subject site.

After the point values for the five ecological parameters are recorded as above, the point values are totaled to achieve a final score for the subject site. The final score is then compared to the 5 point ranges presented below to determine the appropriate response action for the subject site.

Scoring Range	Description of Risk	Response Action
21–25	Definite Risk	Take immediate action (post area, devise control strategy).
16–20	Potential Risk	Consider action as above, but with a lower priority; place under routine tick surveillance.
11–15	Limited Risk	Perform periodic tick surveillance.
6–10	No Present Risk	Consider periodic tick surveillance; no action necessary.
<6	No Risk Likely	No action necessary at this time or at any time in the future.

Cancer Rates after the Three Mile Island Nuclear Accident and Proximity of Residence to the Plant

Maureen C. Hatch, PhD, Sylvan Wallenstein, PhD, Jan Beyea, PhD, Jeri W. Nieves, MS, and Mervyn Susser, MB, BCh

ABSTRACT

Background: In the light of a possible link between stress and cancer promotion or progression, and of previously reported distress in residents near the Three Mile Island (TMI) nuclear power plant, we attempted to evaluate the impact of the March 1979 accident on community cancer rates.

Methods: Proximity of residence to the plant, which related to distress in previous studies, was taken as a possible indicator of accident stress; the postaccident pattern in cancer rates was examined in 69 "study tracts" within a 10-mile radius of TMI, in relation to residential proximity.

Results: A modest association was found between postaccident cancer rates and proximity (OR = 1.4; 95% CI = 1.3, 1.6). After adjusting for a gradient in cancer risk prior to the accident, the odds ratio contrasting those closest to the plant with those living farther out was 1.2 (95% CI = 1.0, 1.4). A postaccident increase in cancer rates near the Three Mile Island plant was notable in 1982, persisted for another year, and then declined. Radiation emissions, as modeled mathematically, did not account for the observed increase.

Conclusion: Interpretation in terms of accident stress is limited by the lack of an individual measure of stress and by uncertainty about whether stress has a biological effect on cancer in humans. An alternative mechanism for the cancer increase near the plant is through changes in care-seeking and diagnostic practice arising from postaccident concern. (*Am J Public Health*. 1991;81:719–724)

Introduction

Shortly after the accident at the Three Mile Island (TMI) nuclear power plant on March 28, 1979, some area residents believed an increase in cancer had occurred, although a detectable excess had not been predicted. In light of the low estimates of radiation emitted, the President's Commission on the Accident at Three Mile Island concluded that the sole health consequence for the population living nearby was mental distress.[1] Because of growing interest in the effects of stress on cancer promotion and progression, we attempted an ecologic test of the relationship between cancer rates and accident stress. Natural disasters have been associated with subsequent cancer risk in previous reports,[2,3] one of which proposed that a similar study be mounted in the Three Mile Island area.[3]

At Three Mile Island, levels of distress among area residents arose immediately after the accident[4–7] and alcohol consumption rose in parallel.[8] While distress levels were elevated, they were generally in the normal range and seemed to abate within nine months, although attitudinal effects such as distrust of authorities and concerns about safety persisted.[4,5] During clean-up and krypton venting at the plant, and restart of the undamaged reactor in October 1985, turmoil at Three Mile Island was ongoing and stress and consequent distress continued or recurred. A year after the accident, a sample of community residents exhibited significantly higher levels of stress hormones and stress symptoms than residents near undamaged nuclear or coal-powered plants although, again, the levels observed were within normal limits[9]; a similar survey five years postaccident found evidence of chronic stress.[10] Periodic resurveys of other cohorts at TMI have also showed persistent distress[11] and even clinical levels of anxiety, depression and hostility, particularly in those living close to the plant.[6,12–14]

In recent years, increased attention has focused on the possibility that stress and other psychosocial factors may affect the carcinogenic process[15–21]—perhaps through the action of corticosteroids on immune function[20,21] or of stress hormones like prolactin on regulation of neoplastic cell growth.[15,19] Indirect mechanisms involving behavioral risk factors also need to be considered, as do possible interactions between stress and other risk factors. While epidemiologic studies have looked for links between psychological states, personality traits, stressful life events and the onset and development of cancer, thus far the evidence is less than conclusive.[22–38] Two recent positive reports are fairly rigorous. In one, survival time for patients with metastatic breast cancer was found to be significantly increased among those randomized to a psychosocial intervention which reduced depression, anxiety, and pain.[39] In the other, severely stressful life events proved to be strongly predictive of a first recurrence of

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Address reprint requests to Maureen C. Hatch, PhD, Division of Epidemiology, Columbia University, School of Public Health, 600 West 168th Street, New York, NY 10032. Coauthors Nieves and Susser are also with that Division; Dr. Wallenstein was with the Division of Biostatistics at Columbia; Dr. Beyea is with the National Audubon Society, New York. This paper, submitted to the Journal March 23, 1990, was revised and accepted for publication January 8, 1991.

Editor's Note: See also related editorial on page 687 of this issue.

breast cancer in carefully matched patients with previously operable tumors.[40]

A biologic rationale connecting stress and cancer does exist Inconsistent results from research in humans may indicate the lack of a true association, or a lack of understanding about how the neuroendocrine and immune systems mediate between psychosocial factors and cancer. In the present report, we summarize an *a priori*, if crude, test of the impact of an objective stress situation on a community's cancer rates.

Methods

The methods of this record-based geographic study are described in detail elsewhere.[41] Briefly, cancers among residents living within 10 miles of the TMI plant in the period 1975–85 were ascertained by review of patient charts at all local and regional referral hospitals (incident cases) and from vital certificates (deaths). The study area was divided into 69 small "study tracts," built up from census blocks, with populations ranging from 500-9,500 (2,300 on average). Cancers were assigned based on residence at diagnosis or at death, and for each study tract, annual rates of cancer incidence and mortality were calculated for the 11 years of the study period.

Cancer Sites Studied

In terms of the accident as a stressful event, we prespecified certain cancers or groupings of cancers for analysis, although there was little literature to guide us. Four categories were selected:

- *leukemias*, chosen because of their responsiveness to glucocorticoids, one of the "stress" hormones[42];
- *lymphomas*, which also respond to glucocorticoids and, in addition, are associated with immunosuppressive states[43];
- *"hormonally dependent" cancers* (breast, endometrium, ovary, prostate and testis), selected because stress hormones can alter concentrations of other hormones that affect cancer cells[19];
- the grouping *"all cancers"*, because in theory stress can influence the biology of all tumors through changes to the immune system.

Proximity as a Surrogate Measure of Stress

In the absence of any direct measure of stress, we used a surrogate measure supported by findings from the earlier psychosocial research at TMI. The measure is based on residential proximity to the Three Mile Island facility. A postaccident evacuation advisory issued by the Governor had focused public attention sharply on the area within a five-mile radius, with evident effect. Not only were evacuation rates higher in the five-mile radius, but so were rates of demoralization as determined by the President's Commission on Behavioral Effects.[4] Another population survey, by the Pennsylvania Department of Health,[5] found that as proximity increased so did the reported levels of perceived threat, of feeling "upset by the accident," and of psychophysiological symptoms. A study of mothers with young children found that those living in a five-mile radius of the TMI plant were more likely to experience a clinical episode of depression or anxiety than mothers living farther away.[14]

To evaluate residential proximity in the study reported here, we assigned each study tract a value corresponding to the distance between its midpoint and the TMI facility. The study area was also divided into three contiguous concentric rings based on natural breakpoints in the distribution of distances.

Proximity as Distinct from Radiation Emissions

While we treat residential proximity as a surrogate for accident-related stress, others have used it to define exposure to radiation near nuclear installations. Factors like wind direction, terrain, and elevation modify the distribution of exposures, however, and in the Three Mile Island area the terrain northeast and northwest of the plant directed emissions away from residents. In estimating radiation exposures, we developed mathematical models that take such modifying influences into account. Overlap between residential proximity and exposure to radiation emissions modeled mathematically is adjusted for in the analysis. A detailed description of the radiation emissions models is given in a separate report.[41]

Statistical Analyses

Logistic regression based on maximum likelihood estimation[44] was used to examine the age-sex-specific probability of cancer in relation to proximity of residence to the plant. Odds ratios derived from the logistic regression estimates were calculated contrasting median distance in the inner and outer rings. Indicator variables were used to adjust for a maximum of 16 age-sex categories (0–14, 15–24,. . . 65–74, 75 +). No adjustment was made for race since the population is over 96 percent White. Using aggregate data for each of the 69 study tracts, we adjusted for urbanization (population per km^2) and social class (median income; percent high school graduates), treating these measures as continuous variables.

The analyses were performed for both the preaccident period (January 1975-March 1979) and the period immediately following the accident (April 1979-December 1985). Where indicated, we used an unconditional logistic regression procedure to control for underlying heterogeneity among study tracts prior to the accident. This analysis, which is a means of adjusting for baseline risk, incorporates a parameter for each study tract, an indicator variable for period (preaccident or postaccident) and a distance variable coded zero before the accident and according to its assigned value after the accident. A conditional analysis based on a Poisson model gave similar results and is not presented here.

For descriptive purposes, we calculated standardized incidence ratios (SIRs) and standardized mortality ratios (SMRs) for the three rings at distances from the TMI facility of 0–6 km, >6–12 km, and >12 km, using age-sex-specific rates for Whites in the national Surveillance Epidemiology End Results (SEER) program (1978–81, Puerto Rico excluded).[45] However, interpretations are based on the regression analyses, which do not use an external standard.

Results

The distribution of population density, income and education by distance from the TMI plant (Table 1) shows that nearby areas are more urban and of lower socioeconomic status; these factors, as well as age and sex, are controlled in the logistic regression models.

Stress in Proximity to the Plant

For all cancers as a group, there is a statistically significant relationship between incidence rates (Table 2) after the accident and residential proximity to the plant (OR = 1.4; 95% CI = 1.3, 1.6). The association is present in both males (OR = 1.4; 95% CI = 1.2, 1.6) and females (OR = 1.5; 95% CI = 1.3, 1.7). Among the specific cancer categories tested, lymphoma shows the strongest association (OR = 1.9; 95% CI = 1.2, 3.0). Leukemias and hormonal tumors show associations of about the same magnitude as for all cancers. In addition, two common tumors

TABLE 1—Descriptive Data for Areas Defined by Distance of Midpoint from the Three Mile Island (TMI) Nuclear Plant

Distance (median) in kilometers	1980 Population[a]	Density per km²	Median Income $	% High School Graduates
0–6 (4.3)	16,115	571	17,662	61.9
>6–12 (10.0)	60,654	316	19,452	68.8
>12 (14.2)	82,915	329	20,192	68.9

a) Midyear population estimates; adjustment has been made to the census figures for April 1, 1980.

TABLE 2—Incidence of Cancer and Proximity to TMI: Standardized Incidence Ratios (SIR)* and Odds Ratios (OR)‡ for Specific Cancer Types before and after the Accident

	Distance from Plant			
Cancer Grouping	0-6km SIR (obs)	>6-12km SIR(obs)	>12km SIR(obs)	OR (95% CI)
All Cancer				
Preaccident	0.9 (166)	0.7 (581)	0.7 (975)	1.2 (1.0,1.4)
Postaccident	1.1 (381)	1.0 (1423)	0.9 (1967)	1.4 (1.3,1.6)
Other Prespecified Sites				
Lymphoma				
Preaccident	1.7 (13)	0.7 (25)	1.0 (51)	1.2 (0.6,2.2)
Postaccident	1.5 (21)	0.9 (54)	0.9 (77)	1.9 (1.2,3.0)
Leukemia				
Preaccident	0.7 (4)	0.5 (12)	0.6 (23)	1.1 (0.4,3.0)
Postaccident	0.9 (9)	1.1 (43)	0.8 (51)	1.4 (0.8,2.5)
Hormonal[a]				
Preaccident	0.9 (45)	0.8 (165)	0.9 (298)	1.1 (0.8,1.4)
Postaccident	1.0 (86)	1.1 (408)	1.0 (581)	1.2 (1.0,1.4)
Selected Common Cancers				
Lung				
Preaccident	0.6 (17)	0.5 (61)	0.6 (116)	1.2 (0.8,1.8)
Postaccident	1.3 (67)	1.0 (214)	0.8 (282)	1.7 (1.3,2.1)
Colon				
Preaccident	0.7 (14)	0.8 (70)	0.8 (104)	1.1 (0.7,1.7)
Postaccident	1.7 (56)	1.3 (185)	1.0 (234)	1.7 (1.3,2.2)

*For calculation of SIRs, expectations are based on SEER (1978–81), age-specific rates for whites (excluding Puerto Rico). Number of observed cancers in parentheses.

‡Odds ratios comparing median distance in the inner ring with median distance in the outer ring, derived from logistic analysis adjusted for age, sex, density, median income, and education.

a) Includes breast, endometrium, ovary, prostate, and testis.

(lung and colon) also show associations with proximity to TMI (OR = 1.7 for each). The observed increase in "all cancers" thus seems to be due to increases in several cancer types.

For the "all cancer" grouping, rates in the area close to the plant are higher before the accident as well as afterwards. We corrected for the preaccident pattern by means of the procedure described above; the odds ratio adjusted for risk at baseline was reduced from 1.4 to 1.2 (95% confidence limits = 1.0, 1.4).

In Figure 1 we show incidence rates for all cancers for each year of the study period by proximity to the plant (three distance rings). By 1982, rates in the area closest to TMI were clearly elevated. The numbers of cases among residents of the inner ring rose from a yearly average of less than 50 to a high of 78 (a value that lies outside the 99% CI around 50 (33.7, 71.3), assuming the number of cases to be a Poisson variable). By 1984, the cancer rate had fallen to preaccident levels and by 1985 was lower than the rates in the more distant rings.

Mortality data (Table 3) suggest no general increase in cancer deaths with proximity to the plant in the postaccident period compared to the preaccident period. Neither do mortality rates for all cancers by year show any evidence of a change (data not presented).

Detection Bias

To test for detection bias arising from postaccident concern about cancer risk, we compared the relationship between stage at diagnosis and residential proximity to the plant in the periods before and after the event. Cancers of the breast, lung, colon, and prostate were selected since screening is common for these sites (see Table 4). With breast cancer, a shift in the postaccident period to an earlier stage at diagnosis is seen only among residents of the middle ring. For prostate cancer, residents of the inner ring actually show a decrease in the proportion of tumors diagnosed at an early stage after the accident compared to before. Lung and colon cancer show the reverse: after the accident, a higher proportion of tumors was diagnosed early among those living close to the plant.

Discussion

In an admittedly crude test of an accident-stress hypothesis, we found an increase in cancer following the accident that related to proximity of residence to the Three Mile Island nuclear plant. Cancer rates in those living nearest TMI rose in 1982, three years postaccident, remained raised for another year, and then declined. The relationship with proximity is unlikely to be explained by better case ascertainment for residents of the inner ring since, in the interests of comprehensive case-finding, we reviewed medical records at all hospitals within a 30-mile radius and at regional referral centers.

Residential proximity to the TMI plant was related to cancer rates prior to the accident as well as after, suggesting the presence of risk factors that were not sufficiently controlled by our adjustments for urbanization and social class. As a means of handling the confounding effects of unmeasured factors (e.g. cigarette smoking), we therefore adjusted for baseline risk. With this adjustment, the odds ratio for residential proximity after the accident was reduced to borderline statistical significance. Nonetheless, the postaccident pattern involving a sudden peak of excess cancers in residents near the plant does appear to be distinct from the preaccident pattern, which showed cancer rates in the inner ring to be slightly higher throughout the period.

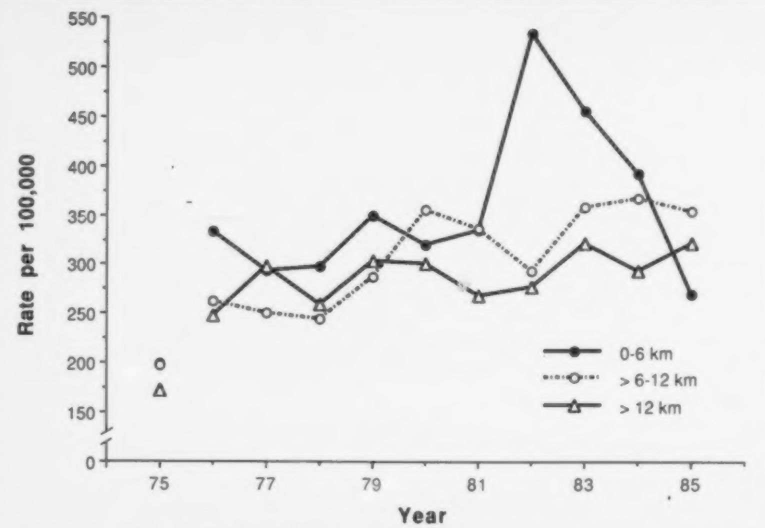

FIGURE 1—Yearly Age-adjusted Incidence Rates (100,000) of all Cancers, 1975-85, for Three Areas Defined by Distance from the TMI Nuclear Plant.

Note: In all three areas, figures were low for 1975, when hospital records were initially computerized.

TABLE 3—Mortality from Cancer and Proximity to TMI: Standarized Mortality Ratios (SMR)[a] and Odds Ratios (OR)[b] for Specific Cancer Types before and after the Accident

Cancer Grouping	Distance from Plant 0-6km SMR(obs)	 >6-12km SMR(obs)	 >12km SMR(obs)	OR (95% CI)
All Cancer				
Preaccident	1.3 (121)	1.0 (396)	1.0 (649)	1.2 (1.0,1.4)
Postaccident	1.0 (156)	1.1 (775)	1.0 (1119)	1.9 (0.9,1.2)
Other Prespecified Sites				
Lymphoma				
Preaccident	2.6 (9)	1.0 (15)	1.1 (28)	1.5 (0.7,3.1)
Postaccident	0.5 (3)	1.0 (27)	0.9 (38)	0.6 (0.3,1.3)
Leukemia				
Preaccident	0.7 (3)	0.6 (12)	1.1 (33)	0.5 (0.3,1.2)
Postaccident	0.8 (6)	1.2 (40)	0.9 (47)	1.0 (0.6,1.9)
Hormonal[c]				
Preaccident	1.2 (21)	1.2 (86)	1.0 (118)	1.4 (0.9,2.0)
Postaccident	0.7 (19)	1.2 (155)	1.2 (229)	0.8 (0.6,1.0)
Selected Common Cancers				
Lung				
Preaccident	1.2 (26)	0.8 (80)	0.9 (135)	1.4 (1.0,2.0)
Postaccident	1.0 (40)	1.0 (164)	0.9 (241)	1.3 (1.0,1.6)
Colon				
Preaccident	0.8 (9)	0.9 (40)	1.2 (84)	0.6 (0.4,1.1)
Postaccident	1.4 (25)	1.0 (80)	1.0 (125)	1.1 (0.8,1.6)

[a] For calculation of SMRs, expectations are based on SEER (1978-81), age-specific rates for whites (excluding Puerto Rico). Number of observed cancers in parentheses.

[b] Odds ratios comparing median distance in the inner ring with median distance in the outer ring, derived from logistic analysis adjusted for age, sex, density, median income, and education.

[c] Includes breast, endometrium, ovary, prostate, and testis.

Why a peak of cancer cases should arise in 1982, three years after the accident is not at all obvious. It could relate to increased cancer surveillance near the plant, particularly as the increase appears to involve several different cancer types. If residents changed their medical care habits following the accident, so that cancer was diagnosed earlier, a "screening effect" could arise, resulting in an initial inflation in cancer rates.[46] We sought and did not find consistent evidence of earlier detection near the plant in our comparison of preaccident and postaccident diagnostic trends. The absence of an increase in mortality rates—particularly for lung cancer which is rapidly fatal—may point to a screening bias, however, as may the dip in the cancer incidence rate near the plant in 1985, the last year of the study period. Distinguishing between improved surveillance, potentially stress-related, and a direct biological effect of stress is difficult with these data, particularly if, as seems likely, stress acts to accelerate the appearance of cancer rather than to initiate new cases.

In terms of a stress hypothesis, the study's major limitation is the lack of any measure of stress at the individual level. Using distance of residence from the plant to indicate stress is bound to introduce measurement error since individuals vary in both their perception of stressful events and response to them. It does seem reasonable to assume, however, that residents living close to the plant experienced more stress on average than residents living farther away. The Governor's evacuation advisory, issued two days after the accident, focused on a five-mile radius, and investigations of possible mental health consequences found that stress responses among area residents were related to distance from the plant.[4,5,9,14] In our study, however, the proximity measure is vulnerable to migration effects: inmigrants will not have been exposed to the stress of the accident but will be considered so in the analysis, and out-migrants may be self-selected for high stress levels. Given these sources of error, if stress does underlie the observed association between proximity to the plant and cancer risk, its effect could be underestimated. On the other hand, ecological associations frequently overestimate the magnitude of effects at the individual level.[47,48] Even if the stress hypothesis holds, the mechanism could be indirect. The stress of the accident could have increased exposure to behavioral risk factors, such as cigarette smoking, but we have no data with which

TABLE 4—Distribution (%) of Selected Cancers by Stage of Diagnosis* and Proximity to the Three Mile Island Plant for the Periods before and after the Accident

Stage at Diagnosis	0-6k pre v. post		>6-12k pre v. post		>12k pre v. post	
Breast Cancer						
Local	39.3	40.4	39.4	51.8	46.0	48.5
Regional	52.6	42.5	31.3	30.7	32.2	36.5
Distant	8.2	17.0	29.1	17.8	21.7	15.0
Prostate Cancer						
Local	62.2	41.7	56.1	59.8	53.3	60.9
Regional	9.0	13.5	20.2	10.7	17.7	15.6
Distant	28.8	44.9	23.7	29.5	29.0	23.5
Lung Cancer						
Local	30.7	36.9	37.5	24.0	23.6	26.7
Regional	23.9	16.9	13.6	26.2	29.4	25.5
Distant	45.4	46.4	48.9	49.8	47.0	47.9
Colon Cancer						
Local	29.7	39.3	30.9	29.1	27.0	30.8
Regional	43.9	29.9	31.7	34.7	35.9	36.2
Distant	26.4	30.8	37.4	36.2	37.1	33.0

*Among those with known stage of diagnosis.

to evaluate this possibility. Whatever the mechanism, a true effect on cancer produced in so short a time could only manifest itself by acting on preexisting cancers.

We carried out statistical analyses to test whether the gradient in cancer risk with residential proximity—present during routine operation of the Three Mile Island plant prior to the accident, as well as after—could possibly be explained by emissions from TMI. Fitting a multivariate regression model with both measures (proximity and plant emissions modeled mathematically) actually strengthens the association of proximity with cancer incidence. Also, the pattern of rise and fall in diagnosed cancers between 1982 and 1984 would not be expected with radiation.

Conclusion

In an analysis examining cancer rates and residential proximity to the Three Mile Island nuclear plant, we observed a modest postaccident increase in cancer near TMI that is unlikely to be explained by radiation emissions. The increase resulted from a small wave of excess cancers in 1982, three years after the 1979 accident. Such a pattern might reflect the impact of accident stress on cancer progression. Our study lacked a direct, individual measure of stress, however, The most plausible alternative explanation is that improved surveillance of cancer near the TMI plant led to the observed increase. Investigations of cancer patterns in communities with future environmental incidents should consider gathering data on stress and stress-related behaviors, including change in medial care-seeking, as well as on exposure to pollutants. ☐

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Child Care Arrangements of American Children: NCHS Report

A recent report from the National Center for Health Statistics profiles the child care arrangements received by American Children. The report, *Child Care Arrangements*, is derived from the National Health Interview Survey on Child Health, conducted by NCHS in 1988.

The survey on child care arrangements found that two thirds of all children 5 years of age and under had been in day care at some time in their lives, with the proportion increasing with age—from 56 percent for those under age 2 to 80 percent for those 4-to-5 years old. In 1988, approximately 13.3 million children were receiving child care, including 50 percent (almost 4 million) of those under the age of 2.

Of all the children receiving child care, 83 percent had mothers who were employed. The most commonly used child care arrangements were nursery schools or preschools (used by 23 percent) and care provided by a nonrelative in a home other than the child's (21 percent). About one fourth of children receiving child care were cared for alone; 35 percent were in groups of two to three children; 20 percent with four to six children; and 23 percent in groups of seven or more.

The report *Child Care Arrangements*, prepared by Deborah A. Dawson, PhD, and Virginia S. Cain, PhD, of the National Institute of Child Health and Human Development, is available from the National Center for Health Statistics, Centers for Disease Control, US Public Health Service, 6525 Belcrest Road, Hyattsville, MD 20782; (301) 436-8500.

Effectiveness of Source Documents for Identifying Fatal Occupational Injuries: A Synthesis of Studies

Nancy Stout, EdD, and Catherine Bell, BGS

ABSTRACT

Background: The complete and accurate identification of fatal occupational injuries among the US work force is an important first step in developing work injury prevention efforts. Numerous sources of information, such as death certificates, Workers' Compensation files, Occupational Safety and Health Administration (OSHA) files, medical examiner records, state health and labor department reports, and various combinations of these, have been used to identify cases of work-related fatal injuries. Recent studies have questioned the effectiveness of these sources for identifying such cases.

Methods: At least 10 studies have used multiple sources to define the universe of fatal work injuries within a state and to determine the capture rates, or proportion of the universe identified, by each source. Results of these studies, which are not all available in published literature, are summarized here in a format that allows researchers to readily compare the ascertainment capabilities of the sources.

Results: The overall average capture rates of sources were as follows: death certificates, 81%; medical examiner records, 61%; Workers' Compensation reports, 57%; and OSHA reports 32%. Variations by state and value added through the use of multiple sources are presented and discussed.

Conclusions: This meta-analysis of 10 state-based studies summarizes the effectiveness of various source documents for capturing cases of fatal occupational injuries to help researchers make informed decisions when designing occupational injury surveillance systems. (*Am J Public Health.* 1991;81:725–728)

Introduction

Identifying all of the fatal occupational injuries in the United States work force is a difficult task that has not been accomplished with precision.[1,2] Recent literature has questioned the accuracy and completeness of ascertainment of such cases by various sources of information.[1,3–6] Death certificates, Workers' Compensation files, Occupational Safety and Health Administration fatality reports, medical examiner records, reports collected by state health or state labor departments, and various combinations of these have been used to identify cases of fatal work injuries. Each of these sources was developed for different purposes, using different definitions and methods. Because there is no universally accepted gold standard against which to compare, particularly at the national level, it is difficult to measure precisely the proportion of all fatal injuries identified by a specific source.

In a smaller geographic area, such as a state, the pool of fatal work injuries can be defined as the combination of unique cases identified through several sources. Capture rates can then be estimated as the proportion of the total pool identified by each source. This method of assessing the completeness of case ascertainment by various sources of fatal work injury data has been employed in at least 10 states. Results of these studies are summarized here to provide an indication of both the overall average capture rates of the sources, and the variation in capture rates by state.

Sources of Data

The source documents used to identify fatal occupational injuries differ in their primary purposes, methods, and target populations. The advantages and limitations of various sources have been discussed elsewhere.[1,3,7] Issues of validity, completeness, accuracy, timeliness, and the usefulness of data elements in these sources are important considerations in fatality surveillance, but are outside the scope of this paper. The sources examined here are briefly described in terms of their characteristics salient to comparing capture rates.

Death certificates and medical examiner records are the only two available sources that have the potential to identify all cases of fatal work injuries in a given geographic area. Death certificates are filed for each death that occurs in the US and follow a standard format that includes an item indicating whether the death was due to an injury sustained at work. The accuracy and consistency of this item is questionable, however, as no explicit national guidelines exist that define "injury at work."[8,9] Motor vehicle crashes and homicides at work appear to be particularly underreported as work injuries on death certificates.[7,8,10,11]

In most states, medical examiner (ME) or coroner investigations are required for deaths that are due to accidents, homicides, suicides, or deaths that occur unexpectedly.[9] In jurisdictions where ME or coroner records are computerized, or at least centralized, they are a valuable source of information.[12,13] In most states,

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From the National Institute for Occupational Safety and Health. Address reprint requests to Nancy Stout, EdD, Division of Safety Research, National Institute for Occupational Safety and Health, 944 Chestnut Ridge Road, Morgantown, WV 26505. This paper, submitted to the Journal September 27, 1990, was revised and accepted for publication February 19, 1991.

TABLE 1—Percent of Fatal Work Injuries Identified by Various Sources: Summary of State-Specific Studies

			Source Documents				
States	Time Period	N	% Death Certificate	% Workers' Compensation	% OSHA Fatality Reports	% Medical Examiner Records	% State Dept Health or Labor
Wisconsin[11]	1976	161	76	57	33		
Maryland[4,a]	1978	148	67	58	42	90	
Michigan[20,b]	1980	195	88				51
Alaska[21]	1980–1985	422	83	47			
Colorado[7]	1982–1984	425	86	66	32		
California[18]	1983	673	90	40	21		
New Jersey[c]	1984–1985	204	84		34	32	
Texas[d]	1986	872	67	70	27		9
Oklahoma[10]	1985–1986	329	78	53	29	82[e]	
Massachusetts[18]	10/86–11/87	86	88	62	41		22
Range (%)			67–90	40–70	21–42	32–90	9–51
Average (%)			81	57	32	68	27

[a] Susan Baker, Johns Hopkins Injury Prevention Center, personal communication, January 1991.
[b] Figures presented here exclude the "natural deaths" from the original study.
[c] Martha Stanbury, New Jersey Department of Health, personal communication, June 1990. New Jersey data exclude intentional injuries.
[d] Texas Department of Health, Occupational Safety and Health Division. Traumatic Occupational Fatalities, 1986. Unpublished report, 1989.
[e] For ME records with unclear work-relatedness, death certificates were reviewed to verify injury at work or reject false positives.

however, both ME and coroner reports exist as fragmented paper files held in various locations throughout a state, they are inconsistent among jurisdictions, and frequently do not indicate the work-relatedness of a fatal injury.[3]

Employers under Occupational Safety and Health Administration (OSHA) jurisdiction are required to report to OSHA all work-related fatalities of employees. Many of these incidents are then investigated and OSHA investigation reports are filed. Not all workers are under OSHA jurisdiction and not all incidents are investigated. Public sector employees, and self-employed workers, for example, are exempt from reporting requirements. Events such as airplane crashes, motor vehicle crashes, and homicides are investigated by other agencies, and are not routinely identified or investigated by OSHA. Recent studies also suggest substantial underreporting of fatalities that do fall under OSHA jurisdiction.[5,7,14–18]

Workers' Compensation claims can provide valuable detail about the nature of fatal injuries and accurate specification of employment. Like OSHA coverage, however, many workers are not covered by Workers' Compensation laws. Moreover, state compensation laws differ substantially so that covered worker populations are not comparable by state, and the national population of covered workers is difficult to define.

In some states, occupational fatalities are reported to the state department of labor or the state department of health. The specific types of employment and fatal events that are reported vary from state to state. For example, in Massachusetts all construction industry employers must report every accident resulting in death. The Massachusetts Department of Labor and Industries maintains records of these and other work-related fatalities reported voluntarily by employers, workers, unions, police, etc.[18]

The State Health Department source of data for the Texas study cited here refers to fatalities reported by employers participating in the Bureau of Labor Statistics (BLS) Annual Survey of Occupational Injuries and Illnesses by Industry. The survey is limited to establishments under OSHA jurisdiction, employing 11 or more workers and, as a sample survey, is not designed to capture all work fatalities.

Differences in capture rates among these sources of fatal work injury data are due in part to variation in types of employments covered by each source. Although several studies have examined capture rates of sources relative to the population covered by law to measure underreporting,[5,7,15,17] the data presented here do not address underreporting of fatal events but rather the potential of various sources to identify the universe of fatal occupational injuries.

Summary of Study Results

In each of 10 studies presented here, multiple data sources were linked, eliminating duplicate cases, to identify a total "pool" of unique fatal occupational injury cases in a state, and the proportion of that pool identified by each separate source. Table 1 summarizes the results of these studies and presents the capture rates for each source, and the range and average capture rates of the combined study results.

All 10 studies used death certificates as one source of case ascertainment and in all but two studies a greater proportion of the total pool of fatal work injuries was identified by death certificates than by any other source. The average capture rate for death certificates over all the studies was 81 percent.

Workers' Compensation files, used as a source of case ascertainment in eight states, had an average capture rate of 57 percent. The wide range (40 to 70 percent) is likely due to the different worker groups covered by the various state compensation laws.

OSHA fatality reports identified an average of 32 percent of cases. The variation in capture rates by state (21 to 42 percent) may be partially due to differences in the OSHA forms examined. In most states, first reports of fatal injuries were used to identify cases. The California[18] and New Jersey[a] studies, however, used OSHA fatality investigation reports. Since not all reported fatalities are inves-

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[a] Martha Stanbury, New Jersey Department of Health, Personal Communication, June 1990.

TABLE 2—Percent of Fatal Work Injuries Identified by Only One Source, by Source: Summary of State-Specific Studies

		Sole Source Documents				
States	Time Period	% Death Certificate	% Workers' Compensation	% OSHA Fatality Reports	% Medical Examiner Records	% State Dept. Health or Labor
Wisconsin[11]	1976	43	24	0		
Maryland[4]	1978	(N/A)	(N/A)	(N/A)	(N/A)	
Michigan[20,a]	1980	49				14
Alaska[21]	1980–85	53	17			
Colorado[7]	1982–84	34	14	0		
California[19]	1983	56	7	1		
New Jersey[b]	1984–85	31		3	12	
Texas[c]	1986	26	27	1		<1
Oklahoma[10]	1985–86	2[d]	12	<1	11[d]	
Massachusetts[18]	10/86–11/87	28	17	0		2
Range (%)		2–56	7–27	0–3	11–12	<1–14
Average (%)		36	17	<1	12	5

[a] Figures presented here exclude the "natural deaths" from the original study.

[b] Martha Stanbury, New Jersey Department of Health, personal communication, June 1990. New Jersey data exclude intentional injuries.

[c] Texas Department of Health, Occupational Safety and Health Division. Traumatic Occupational Fatalities, 1986. Unpublished report, 1989.

[d] For ME records with unclear work-relatedness, death certificates were reviewed to verify injury at work or reject false positives.

tigated, first reports are likely to identify a greater proportion of deaths than investigation reports.

Two of the three studies that included ME records identified more cases from ME records than any other source.[4,10] One study used death certificates to verify injury at work for an unknown proportion of ME records on which work relatedness was not indicated; the capture rate would have been lower if ME records had been reviewed independently.[10] In the third study, ME records identified the smallest proportion of cases.[a] This range in capture rates, from 32 percent to 90 percent illustrates the magnitude of variation by state in the potential for ME records to ascertain cases of fatal work injuries.

Occupational fatalities reported to state health departments and state departments of labor under various legislation are grouped together in Table 1 for simplicity. The average capture rate, (27 percent) is presented as a rough indication of the usefulness of such reports for identifying fatal injury cases. This average, however, is based on only three studies, and substantial differences in reporting requirements limit the comparability of these data by state.

Table 2 presents the proportion of the pool of fatal injury cases identified by only one source of information in each study. The range and average sole source capture rates provide an estimate of the proportion of the universe that might be identified by only one specific source and not the others. Obversely, the average for each source presented in this table may be interpreted as the additional proportion of the total pool of cases that might be identified by using this as an additional source of case ascertainment.

These capture rates are a function of the total pool of fatalities identified in each state, and the number and types of source documents used to define that pool. The sole source averages are presented in Table 2 as crude estimates for the purpose of comparison, rather than as precise means.

Over all the studies, death certificates uniquely identified the greatest proportion of fatal work injuries (36 percent), followed by Workers' Compensation records (17 percent), and ME records (12 percent). State labor department records (5 percent) and OSHA fatality reports (less than 1 percent) contributed relatively few additional cases to the total pool.

Discussion

Generally, if surveillance is limited to one source of information, death certificates will likely ascertain a greater proportion of the universe of fatal work injuries than will the other sources examined here. For national level surveillance, death certificates are unique among these sources in their combined potential to identify all fatal work injuries, and their comparability between states. However, the accuracy and consistency with which the injury at work items on death certificates are completed varies by state. Current efforts by the Agency for Vital Records and Health Statistics (AVRHS), National Institute for Occupational Safety and Health (NIOSH), and the National Center for Health Statistics (NCHS) to increase the validity and reliability of this item could vastly improve our ability to identify the universe of fatal occupational injuries at both state and national levels.

Under some jurisdictions, ME records may provide more complete case ascertainment than other sources. However, they also show the greatest variation by state in their ability to identify cases. Efforts to computerize ME records in a consistent format are currently underway in several states. Uniform automation of ME records could substantially improve fatal work injury surveillance within states and enable interstate comparisons.

The combined results of these studies suggest that Workers' Compensation records are not a useful sole source for case ascertainment of work fatalities. As a supplemental source, however, Workers' Compensation records could be expected to add the greatest proportion of cases to a vital records-based system. OSHA fatality reports have limited utility in fatal injury surveillance. Alone, they might be expected to identify about one-third of all cases and, as an additional source, they might capture an additional 1 percent of cases.

Because the definition of state health department and state labor department fatality reports varies so widely by state, generalized conclusions about this source cannot be drawn. In developing surveillance systems, the state-specific characteristics of this potential data source should be considered.

The combined results of these studies highlight several salient points regarding fatal work injury surveillance. The most prominent of these is that using multiple sources of data will always result in more complete case ascertainment than using one source alone.

The value added by including a subsequent source for case ascertainment depends on which source is added and its characteristics within that geographic area. For example, using state labor department records in addition to death certificates in Michigan improved case ascertainment by 14 percent.[20] In the Massachusetts study, however, where three other sources were also used and where Department of Labor reports had limited coverage, only 2 percent more cases were identified by these records.[18]

In addition to the sources discussed here, other sources should be considered to fill gaps in surveillance systems. For example, all of these sources tend to underreport occupational motor vehicle fatalities and likely undercount homicides. Police report forms and Highway Traffic Administration reports could be valuable in identifying such cases, if efforts were made to include an indication of work-relatedness on these reports.

When designing and developing fatal occupational injury surveillance efforts, cost-benefit analysis and value-added formulas are often considered to determine which sources of data will provide the most effective results given study objective and resource limitations. It is hoped that this synopsis of studies will help researchers make informed decision when designing surveillance systems, recognize limitations of various sources, and support improvements to existing sources so that we might all better understand the magnitude and nature of fatal occupational injuries in the United States. ☐

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Female Homicides in United States Workplaces, 1980–1985

ABSTRACT

Background: Women, while noted for low occupational injury mortality rates, are more likely to die as victims of assault than from any other manner of injury at work.

Methods: From the National Traumatic Occupational Fatality surveillance data, 950 women were identified who were fatally assaulted at work. Homicide rates were calculated for the demographic and employment characteristics of these women. Risk ratios among types of lethal injuries were examined.

Results: During 1980–1985, the crude six-year workplace homicide rate was 4.0 deaths per million working women: one twentieth the homicide rate of the US female population. Decedents ranged from 16 years (the lowest age included in the data base) to 93 years of age. Working women older than 65 years had the highest age-specific homicide rate, 11.3 per million. Women younger than 20 had the lowest, 2.5 per million per year. Homicide rates for women of races other than White were nearly twice as high as those of Whites.

The leading causes of death were gunshot wounds (64 percent), stabbings (19 percent), asphyxiations (7 percent), and blunt force trauma (6 percent). Nearly 43 percent of the deceased women had been employed in retail trade: 8.7 per million employed women annually.

Conclusions: During 1980–1985, only 6 percent of the nation's victims of work-related injury deaths were female: 41 percent of those women were murdered. Homicide is currently the leading manner of traumatic workplace death among women in the United States. (*Am J* Public Health. 1991; 81:729–732)

Catherine A. Bell

Introduction

Injury surveillance conducted by the National Institute for Occupational Safety and Health (NIOSH), Division of Safety Research, has shown that from 1980 through 1985, homicide was the third leading manner of workplace death in the United States, surpassed only by motor vehicle crashes and machine-related fatal injuries. Nearly 13 percent of all occupational injury deaths were homicides. During 1980–85, 47 percent of the United States work force was female,[1] yet only 6 percent of the nearly 7,000 annual fatal work-related injuries were suffered by women: 41 percent of these women were homicide victims.[2]

Two state-based studies have shown that homicide is a prominent form of lethal occupational injury among women. In a study of fatal work-related injuries in Texas, 1975–80, 53 percent of the 348 women were murdered.[5] Only 13 percent of the fatally injured men were homicide victims.[6] The occupational homicide rate for women was one-third that for men: 7 per million employed women per year compared to 21 per million employed men. In California during 1979–81, the rate of work-related homicide among employed women was 5 per million compared to 22 per million among men.[7] Both studies showed that while a woman was at lower risk of being murdered at work than a man, a greater proportion of female traumatic occupational deaths were due to homicide.

Recently, NIOSH began using demographic and injury event information from state death certificates to examine US fatal occupation injuries, including homicides. This report describes the female victims of work-related homicide during 1980–85, as reported to the National Traumatic Occupational Fatality (NTOF) surveillance system.

Methods

The NTOF data base is composed of death certificates from all state vital statistics reporting units in the US that meet the following criteria: 1) age at death—16 years or older; 2) a positive response to the "Injury at Work" item; and 3) an external cause of death (International Classification of Diseases, Ninth Revision, E800-E999). The methods and limitations that pertain to NTOF injury surveillance are described fully elsewhere.[8]

All NTOF records of female deaths from 1980 through 1985 that reported homicide as the manner of death were selected for this analysis. Homicide, death due to injuries purposely inflicted by another person, was identified using the element in Part II of the death certificate that queries "Accident, Suicide, Homicide, or Pending" because not all NTOF records had coded cause of death information at the beginning of the study. No effort was made to exclude deaths due to legal intervention. Recording practices in Louisiana, Nebraska, New York, and Oklahoma prevented identification of work-related

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From the National Institute for Occupational Safety and Health in Morgantown, WV, where the author was with the Division of Safety Research at the time of the study. Address reprint requests to Chief, Injury Surveillance Section, Division of Safety Research, National Institute for Occupational Safety and Health, S-109, 944 Chestnut Ridge Road, Morgantown, WV 26505. This paper, submitted to the Journal February 20, 1990, was revised and accepted for publication February 19, 1991.

TABLE 1—Female Occupational Homicides in the United States[a] by Year, 1980–85: Numbers and Annual Rates[b]

Year	Number	Deaths per 1,000,000
1980	173	4.75
1981	163	4.36
1982	171	4.44
1983	143	3.65
1984	147	3.59
1985	150	3.57
All Years	950	4.04

[a] Excluding deaths from Louisiana, Nebraska, New York, and Oklahoma.
[b] Deaths per 1,000,000 workers.

TABLE 2—Female Occupational Homicides in the United States[a] by Age, 1980–85: Numbers and Six-year Rates[b]

Age (years)	Number	Deaths per 1,000,000
16–19	43	2.50
20–24	155	4.47
25–34	286	4.34
35–44	159	3.17
45–54	138	3.81
55–64	101	4.11
65+	68	11.28

[a] Excluding deaths from Louisiana, Nebraska, New York, and Oklahoma.
[b] Deaths per 1,000,000 workers.

intentional injuries over the study period, so homicide data from these four states were not available.

Employment information was coded from the "usual" industry and occupation entries on the certificates. The Standard Industrial Classification[9] system was used to categorize usual industry of employment. Occupation information was coded into major groups with the 1970 Bureau of Census Classifications for deaths in 1980 through 1982 and the 1980 Bureau of Census Classifications for deaths in 1983 through 1985. The two coding schemes ensured agreement between incidence and underlying employment information because the source of the population data used in the rate calculations changed coding schemes in 1983.[10,11]

To calculate mortality rates, the Annual Summaries of EMPLOYMENT AND EARNINGS[12] provided denominator data for the US. The GEOGRAPHIC PROFILE OF EMPLOYMENT AND UNEMPLOYMENT[13] was used to correct the denominator data used in age-, race-, and employment-specific rates by excluding the estimated work force of the four states which did not provide homicide data to NTOF (nearly 10 percent of the female work force of the US).

Homicide frequencies and annual and six-year mortality rates were calculated for demographic and employment characteristics. Victims whose occupations or industries of employment were unclassifiable were not included in employment-specific rate calculations, although they were included in all other rates. Age- and race-specific rates were estimated. Race-specific rates were calculated for Whites (including Hispanics), Blacks, and Others (Asians and American Indians). Definitive employment data were not available for Hispanics, Asians, or American Indians. Risk ratios (RR) contrasting types of injury (e.g. firearms versus not firearm) within racial subsets were estimated.

Deaths assumed to involve close physical contact, such as stabbing or beating, were compared to those resulting from gunshot wounds that could be inflicted from a distance. Two categories of close physical contact deaths were defined: multiple stabbing and a broader category that included all rape, stabbing, slashing, blunt force, and manual asphyxiation deaths. These were contrasted to the gunshot wounds through risk ratio estimates. Ninety-five percent Taylor Series confidence intervals were constructed around all risk ratio estimates.[14]

Results

NTOF data described 950 women who were murdered at work during 1980–85 (Table 1). The women's ages ranged from 16, the youngest documented in NTOF data, to 93 years. The six-year age-adjusted homicide rate was 4.5 deaths per million working women. The highest homicide rate, 11.3 per million, was among working women 65 years of age and older (Table 2).

Black women were 1.8 times more likely to be killed than Whites. The RR for homicide among Blacks contrasted to all other women was 1.7 (95% CI: 1.4, 2.0) and for the Other racial category was 1.9 (95 % CI: 1.4, 2.5). Hour of injury was reported for 680 cases: homicides were most frequent between 4:00 and 5:00 pm. However, 30 percent of the women were murdered between 6:00 pm and midnight and 69 percent died between 3:00 pm and 7:00 am.

Industry of Employment

Nearly 43 percent of the homicide victims were employed in the Retail Trade Industry which also had the highest six-year homicide rate, 8.7 per million retail employees. Forty-one percent of these women worked in food stores, and 34 percent in eating and drinking establishments.

The Transportation, Communication, and Public Utilities Industry had the second highest rate, 4.0 per million: 34 percent were employed in motor freight transportation and warehousing and 18 percent in local passenger transportation with buses and taxis. While 23 percent of the women worked in the Service Industry, their homicide rate was quite low, 2.22 per million: 22 percent of these women worked in education, 16 percent in health services, and 13 percent in hotels, motels, and other lodging.

In Public Administration, 33 percent of the women were employed in justice, public order, and safety, 28 percent in general government, and 23 percent in the military. Industry of employment was not coded for 14 percent of the decedents, due to lack of specific employment information on the death certificates (Table 3).

Occupation

Despite the changes in the occupation classification schemes in 1983, 71 percent of the murdered women were employed in only four occupational groups: sales personnel (n = 184), clericals (n = 169), service employees (n = 182), and supervisors (executives, managers, and administrators) (n = 151). During 1980-82, retail store clerks and cashiers, classified as clericals, accounted for 26 percent of the decedents' occupations: their homicide rate was only 3.4 per million. Sales workers, such as sales clerks and owners of retail establishments, died at a rate of 8.3 per million. During the next three years, retail clerks and cashiers were classified as sales occupations. These accounted for 27 percent of the 1983–85 deaths with a three-year rate of 7.6 per million.

For the six-year period, 105 of the 175 women in service occupations were employed in eating and drinking establishments. At least 52 of these worked in bars, taverns, or lounges. Nineteen women were employed as law enforcement officers or security guards. Lack of national

TABLE 3—Female Occupational Homicides in the United States[a] by Industry, 1980–85: Numbers and Six-year Rates[b]

Industry	Number	Deaths per 1,000,000
Agriculture/Forestry/Fishing	4	1.11
Construction	6	2.15
Manufacturing	44	1.23
Transportation/Communication/Public Utilities	38	3.97
Wholesale Trade	9	1.52
Retail Trade	405	8.66
Finance/Insurance/Real Estate	55	2.78
Services	219	2.20
Public Administration	40	3.92
Not Classified	130	—

[a] Excluding deaths from Louisiana, Nebraska, New York, and Oklahoma.
[b] Deaths per 1,000,000 workers.

estimates for employment in these specific occupations made rate calculations impossible.

Employment as a supervisor increased a woman's risk of homicide. Decedents most commonly managed restaurants, apartment complexes, or retail establishments. Ten percent of the women owned their own businesses.

Region

Only 24 (3 percent) of the women were murdered outside the states where they resided. Despite the lack of reporting from Louisiana and Oklahoma, over half (52 percent) of the murders took place in the South at a rate of 5.7 per million women working in the reporting states. The Western states had the second highest rate, 4.1 per million, and 23 percent of the homicides. The North Central states, where Nebraska deaths were not reported, had a homicide rate of 2.5 per million and 18 percent of the deaths. Only 7 percent of the fatalities were reported from the Northeast at a rate of 1.2 per million. This region's figures are seriously compromised by the omission of New York deaths from the data. Of the 46 states reporting, rates indicated that women were at highest risk of being murdered at work in Nevada, Mississippi, the District of Columbia, Florida, and Texas.

Method of Homicide

Firearm injuries caused 609 deaths (64 percent): 148 (16 percent) noted multiple gunshot wounds. Stabbing and slashing caused 181 deaths (19 percent). Multiple stabbing was mentioned in 103 cases (11 percent). Decedents were asphyxiated in 69 (7 percent) or died from blunt force injuries in 57 (6 percent) of the cases. Together, fires, explosions, motor vehicle impacts, poisoning, and sexual assault were reported in fewer than 4 percent of all cases.

Lethal gunshot wounds were most frequent in the 25–54 year age range. The firearm homicide rate among Black females averaged 4.4 per million compared to 2.2 per million among Whites. Black women were slightly more likely to die from gunshot than other wounds (RR, 1.1, 95% CI: 1.0, 1.2). The risk of Hispanic women dying from firearm relative to other injuries (RR) was 1.4 (95% CI: 1.2, 1.6).

Multiple stabbings were prominent, but not significantly elevated, among the oldest women. However, the risk of an elderly woman dying from any form of physical contact relative to gunshot wounds (RR) was 3.6 (95% CI: 2.6, 4.9). Asian and American Indian women were slightly more likely to be killed in trauma involving physical contact than from gunshot wounds (RR of 1.4, 95% CI: 1.0, 2.0), while Black women were more likely to die from gunshot wounds (RR = 0.7, 95% CI: 0.5, 0.9).

Discussion

While the limitations of death certificate information in general[15–18] and NTOF in particular[8] have been well defined, a few merit discussion in terms of occupational homicide surveillance. Surveillance based on death certificates alone underreports cases of fatal occupational injuries. State studies have shown between 67 percent and 88 percent of all traumatic workplace deaths can be identified using death certificates alone.[19] Legal sources such as police files and medical examiner records are more detailed; however, they do not regularly identify work-related cases and are not readily accessible in most US jurisdictions. Death certificates provide systematic accounts of each traumatic death in the US, and not those reported to have occurred at work. Death certificates record "usual" occupation and industry of employment, not necessarily employment at the time of injury. Studies show that between 64 percent and 74 percent of reported usual occupations correspond with decedents' most recent occupations, and between 60 percent and 76 percent of the usual industry entries correspond with most recent industries of employment.[16-18]

Both formal state policies and individual practices affect case ascertainment and data accuracy. The most notable weakness is lack of standard criteria for completing the injury at work item; it is not rigorously defined and is subject to interpretation. For example, in Louisiana a state policy existed where "injury" was interpreted as "accidental injury." Until 1985, the injury at work item was not completed for intentional injuries. Even when injuries that take place at work are correctly recorded, some states automate that information only for unintentional injuries. In these states all homicides and suicides must be screened to find each noted as having occurred at work. This screening is tedious and through human error intentional injury cases may be missed.

A certifier may also impose an idiosyncratic definition of work that excludes or obscures employment. No prostitutes or illicit drug dealers are identifiable in the NTOF data. Juveniles are rarely considered "at work." Deaths perpetrated by acquaintances or family members may be screened by a certifier who considers "at work" to mean solely work-related. Homicides are also misclassified as accidents, suicides, or deaths of undetermined cause. For such reasons, the cases here underreport victimization among working women in the US.

While such factors can hinder hazard identification and ultimately injury control, NTOF surveillance rapidly identified homicide as the major occupational hazard for the nation's women. Homicide risk may be associated with exposure to the public, the exchange of money, and evening employment.[5–7,20] Women employed in the retail industry, particularly in food stores, were at greatest risk. Many owned or managed groceries, taverns, and restaurants. Elevated homicide rates in motor freight and warehousing, local passenger transportation, government, and law enforcement call for heightened safety for workers in these jobs. Detailed studies of homicide risk in education, health services, hotels, motels, and other lodging places are also warranted.

The role of victim age in these deaths is still not clear. Older women may be attractive targets and/or less likely to survive traumatic assault. This analysis also showed that older women were often self-employed or small-business owners. While Davis, *et al*, argued that high workplace homicide rates in Texas women age 65 years or older did not result from underenumerated employment,[5] the person-year denominators used in the NTOF age rates may distort the risk estimates. For example, low rates among young women may reflect part-time work. Age-specific homicide rates should be examined using hour-based employment information when available.

Most of the women died from gunshot wounds. Close physical contact deaths were significantly elevated among older White women and among Asians, a group generally noted for low risk of violent death.[21] Because control strategies may depend on discrimination between robbery assault and personal or sexual violence, more detailed studies are needed on the circumstances of workplace homicides.

At least one major convenience store chain has implemented robbery deterrence programs which couple behavioral and environmental components.[22] In evaluated programs, employees have been trained in techniques of conflict resolution and nonviolent response.[21–24] Locked drop-safes, well-lit parking lots, and a work area openly visible to the public are environmental components that appear to reduce robbery and associated mortality.[22]

Convenience stores are essentially small businesses with control strategies that might prove useful in small, owner-operated businesses. Before recommending use of their control strategies developed in other environments, performance in the new environment should be rigorously evaluated.

Hazardous industries may also benefit from regulation, particularly where employers have not reduced homicide risk. The Occupational Safety and Health Administration (OSHA) standards do not yet explicitly address safety from workplace violence, but an increasing number of local ordinances mandate hours for store operation or the presence of multiple employees. The next step in reducing the risk of workplace homicide is to evaluate the effectiveness in various work settings of existing as well as new control strategies. □

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Occupation, Industry, and Fatal Motor Vehicle Crashes in 20 States, 1986–1987

ABSTRACT

Background: Motor vehicles are a leading cause of injury on and off the job.

Methods: To describe the occurrence of fatal motor vehicle crashes in the working-age population, a case-control study was conducted among persons 15–64 years old who died in 1986 or 1987 in any of 20 states reporting death certificate occupational data to the National Center for Health Statistics. Cases were occupants of motor vehicles (excluding motorcycles) who died of injuries sustained in a crash (E810-E825, 9th revision International Classification of Diseases), and controls died of any other cause.

Results: Transportation-related occupations had more than the expected number of deaths for men (odds ratio = 1.6, 95% confidence interval = 1.4 – 1.9) and women (odds ratio = 2.1, 95% confidence interval = 1.0 – 4.5), as did managerial occupations (odds ratio = 1.3 and 1.7, 95% confidence interval = 1.1 – 1.5 and 1.3 – 2.2 for men and women, respectively). Men in mining and oil drilling occupations and the wholesale trade industry also had excess crash deaths. Among women, additional excesses were observed among professional specialists and in several groups with few deaths, including agriculture, construction, and the military, whereas homemakers had markedly fewer vehicle crash deaths than expected.

Conclusions: These findings encourage further investigation of this important safety hazard, but also call attention to a need for more detailed studies and improved data. (*Am J Public Health.* 1991;81:733–735)

Dana P. Loomis, MSPH, PhD

Introduction

Motor vehicles cause more fatal occupational injuries in the United States than any other agent.[1] Previous studies providing data on this problem have focussed on single states or selected occupational groups,[2-6] but this investigation, using death certificate data from the National Center for Health Statistics, was undertaken to provide a description of the overall epidemiological pattern of motor vehicle-related fatalities in the United States working population to be used in planning future research and prevention efforts.

Methods

The subjects of this study were residents of 20 states (Alaska, Colorado, Georgia, Indiana, Kansas, Kentucky, Maine, Missouri, Nevada, New Hampshire, New Mexico, North Carolina, Ohio, Oklahoma, Rhode Island, South Carolina, Tennessee, Utah, Vermont, and Wisconsin) who died in 1986 or 1987. Only these states coded occupation and industry data from death certificates and reported them to the National Center for Health Statistics (NCHS) in at least one of those two years. Occupation and industry coding are done by the states according to the 1980 US Census Bureau occupation and industry classification system.[7]

Subjects were selected from 1986 and 1987 mortality data tapes from NCHS. All 5,031 men and 1,776 women 15–64 years old who died of injuries received as drivers or passengers in crashes of motor vehicles other than motorcycles (coded E810-E825 with 4th digit equal to 0 or 1 according to the International Classification of Diseases, 9th Revision), and who had an occupation other than "student," "volunteer," "never worked," or "retired" on the death certificate were selected as cases. A control group of 7,167 men and 4,369 women was formed by selecting a 5 percent simple random sample of deaths from all causes other than motor vehicle crashes from the same NCHS tapes using restrictions identical to those applied to the cases.

Standard case-control odds ratios were computed by treating each occupational category in succession as the "exposed" group, and all other categories combined as "unexposed."[8] The odds ratios can be interpreted as the ratio of the observed number of motor vehicle crash fatalities in each occupational category to that expected based on the experience of all others.[9] Age-adjusted odds ratios were estimated by the Mantel-Haenszel method, and test-based confidence intervals were computed for both the crude and adjusted estimates.[8] Because female homemakers were the largest single occupational group, they were considered separately using all women with employment outside the home as the unexposed referent, but were excluded from all other analyses involving women.

Results

Occupation

Crude and age-adjusted odds ratios for motor vehicle crash fatalities by occupation are shown in Table 1. Without adjustment for age, men in white collar oc-

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Address reprint requests to Dana P. Loomis, PhD, Department of Epidemiology, School of Public Health, CB 7400, University of North Carolina, Chapel Hill, NC 27599-7400. This paper, submitted to the Journal August 31, 1990, was revised and accepted for publication February 19, 1991.

TABLE 1—Odds Ratios (95% confidence intervals in parentheses) for 1986–87 Motor Vehicle Crash Deaths for Men and Women Aged 15–64 years by US Census Bureau Occupational Category (unexposed referent group is all other occupations)

Occupation	Men				Women[a]			
	Cases	Controls	Crude	Age-adjusted	Cases	Controls	Crude	Age-adjusted
Professional specialists	243	501	0.7 (0.6, 0.8)	0.9 (0.8, 1.1)	175	310	0.9 (0.8, 1.1)	1.2 (1.0, 1.5)
Technicians & related support workers	99	104	1.4 (1.0, 1.8)	1.1 (0.8, 1.5)	58	74	1.3 (0.9, 1.9)	1.2 (0.8, 1.8)
Sales	403	605	0.9 (0.8, 1.1)	1.2 (1.0, 1.4)	164	246	1.0 (0.8, 1.3)	0.9 (0.7, 1.2)
Administrative support & clerical	158	255	0.9 (0.7, 1.1)	0.9 (0.7, 1.1)	269	503	0.9 (0.7, 1.0)	0.8 (0.6, 0.9)
Services	369	597	0.9 (0.8, 1.0)	0.7 (0.6, 0.8)	314	592	0.9 (0.7, 1.0)	0.8 (0.6, 1.0)
Farming & agriculture occupations	238	391	0.9 (0.7, 1.0)	1.0 (0.8, 1.2)	19	27	1.2 (0.7, 2.1)	1.4 (0.6, 3.0)
Forestry, fishing & hunting occupations	48	54	1.3 (0.9, 1.9)	1.0 (0.7, 1.6)	0	1	0 —	0 —
Mechanics & repairers	351	422	1.2 (1.0, 1.4)	1.1 (0.9, 1.4)	6	10	1.0 (0.4, 2.7)	0.5 (0.1, 1.8)
Construction trades	571	701	1.2 (1.1, 1.3)	1.0 (0.9, 1.2)	9	2	7.5 (2.0, 27.6)	5.7 (1.2, 26.7)
Extractive occupations	56	58	1.4 (1.0, 2.0)	1.8 (1.2, 2.8)	1	1	1.7 (0.1, 25.8)	1.0 (0.0, 855.0)
Precision production workers	280	452	0.8 (0.7, 1.0)	1.0 (0.9, 1.3)	27	54	0.8 (0.5, 1.3)	1.0 (0.5, 1.8)
Machine operators & tenders	334	503	0.9 (0.8, 1.1)	0.9 (0.7, 1.0)	125	195	1.1 (0.9, 1.4)	1.1 (0.8, 1.5)
Fabricators, assemblers & hand workers	146	174	1.2 (1.0, 1.5)	1.0 (0.8, 1.4)	26	55	0.8 (0.5, 1.3)	0.7 (0.4, 1.4)
Inspectors, testers, samplers & weighers	30	40	1.1 (0.7, 1.7)	1.4 (0.7, 2.6)	15	27	0.9 (0.5, 1.7)	1.5 (0.8, 2.8)
Transportation & material moving occupations	537	658	1.2 (1.1, 1.3)	1.6 (1.4, 1.9)	19	16	2.0 (1.0, 3.8)	2.1 (1.0, 4.5)
Handlers, cleaners, helpers & laborers	683	826	1.2 (1.1, 1.3)	0.7 (0.6, 0.8)	80	113	1.2 (0.9, 1.6)	1.0 (0.7, 1.4)
Military occupations	142	203	1.0 (0.8, 1.2)	0.9 (0.7, 1.3)	10	2	8.3 (2.3, 29.7)	1.5 (0.2, 13.4)
Homemakers	4	5	1.1 (0.3, 4.2)	0.3 (0.0, 1.5)	316	1460	0.3 (0.2, 0.3)	0.4 (0.4, 0.5)

[a] Among women the referent group for each occupation in the paid workforce is all other such occupations, i.e. homemakers are excluded; for female homemakers, the referent group is all occupations except homemaker.

TABLE 2—Odds Ratios (95% confidence intervals in parentheses) for 1986–87 Motor Vehicle Crash Deaths for Men and Women Aged 15–64 years by US Census Bureau Industry Category (unexposed referent group is all other industries)

Industry	Men				Women[a]			
	Cases	Controls	Crude	Age-adjusted	Cases	Controls	Crude	Age-adjusted
Mining	96	107	1.3 (1.0, 1.7)	1.3 (0.9, 1.9)	1	4	0.4 (0.1, 3.5)	0.1 (0.0, 1.1)
Construction	919	1075	1.3 (1.2, 1.4)	1.0 (0.9, 1.1)	29	16	3.0 (1.7, 5.5)	3.2 (1.4, 7.4)
Manufacturing	1249	1880	0.9 (0.9, 1.0)	1.0 (0.9, 1.1)	303	482	1.1 (0.9, 1.2)	1.1 (0.9, 1.4)
Transportation, communication & utilities	499	705	1.0 (0.9, 1.1)	1.4 (1.2, 1.6)	74	76	1.7 (1.2, 2.3)	2.0 (1.3, 3.0)
Wholesale trade	165	197	1.2 (1.0, 1.5)	1.4 (1.1, 1.8)	24	40	1.0 (0.6, 1.7)	0.8 (0.4, 1.5)
Retail trade	603	685	1.3 (1.2, 1.5)	1.0 (0.8, 1.1)	340	437	1.4 (1.2, 1.6)	0.9 (0.7, 1.1)
Finance, insurance & real estate	107	206	0.7 (0.6, 0.9)	1.0 (0.8, 1.3)	92	128	1.2 (0.9, 1.6)	1.2 (0.8, 1.7)
Business & repair services	320	391	1.2 (1.0, 1.4)	1.0 (0.8, 1.2)	38	55	1.2 (0.8, 1.8)	0.9 (0.5, 1.6)
Personal services	80	149	0.8 (0.6, 1.0)	0.7 (0.5, 0.9)	93	262	0.6 (0.4, 0.7)	0.7 (0.5, 1.0)
Entertainment & recreation services	46	68	1.0 (0.7, 1.4)	0.7 (0.5, 1.2)	14	22	1.1 (0.5, 2.1)	1.1 (0.4, 2.9)
Professional & related services	249	534	0.7 (0.6, 0.8)	0.9 (0.7, 1.0)	338	673	0.8 (0.7, 0.9)	0.9 (0.8, 1.1)
Public administration	180	347	0.7 (0.6, 0.9)	1.0 (0.8, 1.3)	49	138	0.6 (0.4, 0.8)	0.6 (0.4, 1.0)
Military	147	209	1.0 (0.8, 1.2)	1.0 (0.7, 1.3)	9	2	7.5 (2.0, 27.6)	1.3 (0.1, 12.5)

[a] Excludes homemakers.

cupations, except for technicians and related workers, tended to have essentially the expected number of crash deaths or somewhat fewer, whereas most blue collar occupations had a small excess of deaths. With age adjustment, "extractive" occupations (miners and oil well drillers), and transportation and material moving occupations had the highest odds ratios, with 1.8 and 1.6 times more deaths than expected, respectively. Executives, administrators, managers, and sales workers had more modest excesses of deaths.

For women employed outside the home, some of the largest excesses of vehicle crash fatalities occurred in occupations with relatively few deaths, notably construction trades, transportation, agriculture, and the military. Odds ratios were also elevated, although to a generally lesser degree, in more populous categories, including managerial occupations, professional specialists, and technicians. Female homemakers had very low crash mortality compared to women with paid employment.

Industry

Odds ratios indicating the association of motor vehicle crash deaths with Census Bureau industry categories are shown in Table 2. Among men, most industries had essentially the expected number of crash deaths. However, the categories of transportation, communication, and utilities and wholesale trade had 40 percent more deaths than expected when adjusted for age.

Construction, followed by transportation, communication, and utilities, were the industrial categories with the largest excesses of crash fatalities among women employed outside the home, with about three and two times the expected number of cases, respectively. The more modest excess of deaths among women in the finance, insurance, and real estate category may be more quantitatively important given the larger number of deaths in that group. Public administration and personal services are populous categories with substantially fewer deaths than expected. Crude and adjusted odds ratios tended to be similar except for the military category, for which the odds ratio declined sharply with adjustment for age.

Discussion

This 20-state mortality study describes patterns of motor vehicle crash fatality for men and women by occupation and industry. Such information has not been available before for a geographically diverse population including a wide spectrum of occupations. The vital statistics data used in this study offer several advantages for identifying groups of workers at increased risk: death registration is complete, comparability of data between cases and noncases is assured, and large studies including many states and all occupational groups can be done at minimal cost. On the other hand, studies of this type share generic limitations of design and data quality which have been discussed elsewhere.[9-11]

The major interpretational challenge of this study is that the data do not allow injuries which occurred while the decedent was working to be distinguished from those which occurred during other activities. To minimize the proportion of nonworkers in the study, only individuals ages 15 to 64 years with an identifiable occupation indicated on their death certificate were included. However, deaths due to injuries which did not occur at work were also necessarily included, because the states do not report to NCHS the death certificate field that identifies on-the-job injuries.[10] A field coded for the place of injury (for example, industrial place, home, or farm) is available from NCHS public data, but it does not appear to effectively distinguish occupational injuries.[12]

Although it would be helpful to separate injuries which occur on the job from those taking place elsewhere, the ability to do so is not central to all potential uses of these data. The present analysis based on occupational groups also approximates the socioeconomic pattern of motor vehicle crash mortality. In addition, the data are relevant to injury prevention efforts, because some programs directed toward workers, like the Occupational Safety and Health Administration's proposed workplace driver training program,[13] have the potential to reduce injury risk both on and off the job. Accurate identification of mortality excesses due to on-the-job hazards remains important for more specific interventions, such as equipment modifications, however.

In spite of some disadvantages, the data used in this study appear to be sensitive enough to capture previously-observed excess vehicle crash mortality among transportation workers.[2,6] This should lend some confidence to the interpretation of the occupational patterns of mortality which emerge from these data. Comparison of the crude and age-adjusted odds ratios may also be useful for interpreting the variation in mortality among different groups of workers: the adjusted odds ratios should most accurately reflect the inherent hazardousness of a job, while the crude measures are also sensitive to the sociodemographic composition of the workforce.

Because of the large number of injuries and deaths due to the use of motor vehicles in work, these hazards should be a priority for injury research and prevention. This study suggests that further investigation of managers and other seemingly low-risk groups, in addition to transportation workers, might be warranted. However, the availability of appropriate data remains an obstacle. Uniform coding of the death certificate indication of injury at work and its inclusion in NCHS public mortality data would improve future surveys based on those data, but studies using more detailed information on individuals and vehicles used for work are also needed to identify factors which could be modified to prevent motor vehicle injuries to workers. □

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An Analysis of Occupational Blood Lead Trends in Manitoba, 1979 Through 1987

Annalee Yassi, MD, MSc, Mary Cheang, MMath, Milton Tenenbein, MD, Geoffrey Bawden, Jerry Spiegel, MA, MSc, and Ted Redekop, MD

ABSTRACT

Background: While regulations for workplace lead exposure become more strict, their effectiveness in decreasing blood lead concentrations and the method by which this is attained have not been evaluated.

Methods: An analysis was conducted of 10,190 blood lead samples from employees of 10 high-risk workplaces collected in Manitoba, 1979–87, as part of regulated occupational surveillance.

Results: A significant decrease in blood lead concentrations was observed overall as well as for each individual company. A 1979 government regulation to reduce blood lead to below 3.38 μmol/L (70 μg/dl) was followed by a drop in blood lead concentrations; a 1983 order to reduce blood leads to below 2.90 μmol/L (60 μg/dl) was not followed by such a drop. Longitudinal analysis by individual workers suggested that companies were complying by use of administrative control, *i.e.*, removing workers to lower lead areas until blood lead levels had fallen, then returning them to high lead areas.

Conclusion: Focusing upon blood lead as the sole criterion for compliance is not effective; regulations must specifically require environmental monitoring and controls. Biological surveillance serves as "back-up" to environmental surveillance and this database illustrates the usefulness of a comprehensive centralized surveillance system. (*Am J Public Health*. 1991;81:736–740)

Introduction

Although lead poisoning has been documented since antiquity, occupational lead intoxication continues to occur.[1–3] Moreover, there is growing consensus that lead causes toxic injury to humans at levels of exposure that only a decade ago were thought to be safe. As Landrigan notes,[1] these subclinical changes represent truly harmful outcomes and are not merely homeostatic or physiologic "adjustments" to the presence of lead.

The US Occupational Safety and Health Administration (OSHA) standard, since 1978, has required the removal from exposure of workers who have confirmed blood lead levels of 2.9 μmol/L (60 μg/dl) or higher, an average of 2.4 μmol/L (50 μg/dl) or higher on three tests, or a medical condition contraindicating continued exposure to lead. Once removed, workers cannot return to a job where they will be exposed to lead until the blood lead level is 1.93 μmol/L (40 μg/dl) or less or the employee no longer has a medical contraindication to lead exposure.[2] Environmental monitoring is also required, to ensure a permissible exposure limit of 50 μg/m.[3]

Evidence is lacking as to the effectiveness of any regulation with respect to lead poisoning. While some states maintain heavy metal registries, these only required the reporting of blood lead concentrations above 1.93 μmol/L (40 μg/dl) (until recently when some states dropped their required reporting level to 1.21 μmol/L [25 μg/dl]). Hence, longitudinal follow up of individuals, as well as by workplace, is not possible.

Manitoba, a Canadian province with a population of one million, has a list of high-risk workplaces that are under government surveillance, as well as a centralized provincial laboratory where all occupational blood lead samples are analyzed. With over 11,000 blood lead concentrations recorded in this data base for more than a decade, linked by individual as well as by workplace, it is possible to evaluate trends in the province overall, as well as by individual company. It is also possible to assess concurrent turnover or rotating of staff (administrative controls). Utilizing this blood lead data base, we address specifically the following questions:

- What is the extent of occupational lead intoxication?
- Have mean blood lead concentrations decreased following government improvement orders and tighter standards?
- What has been the role of administrative measures and engineering controls in any such decrease?

Methods

Lead Measurements

A total of 10,190 blood lead tests were performed as a result of regulatory sur-

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Address reprint requests to Annalee Yassi, MD, MSc, Director, Occupational and Environmental Health Program, Department of Community Health Sciences, University of Manitoba, S113-700 Bannatyne Avenue, Winnipeg, Manitoba, R3E OW3. Coauthor affiliations: At the University of Manitoba, Ms. Cheang is with the Biostatistical Consulting Unit, Dr. Tenenbein is Associate Professor, Departments of Pediatrics, Pharmacology, and Community Health Sciences, and Dr. Redekop is Community Medicine Resident, as well as Medical Officer of Health with Manitoba Health; in the Manitoba Department of Labor, Mr. Bawden is Director of Field Operations, Division of Workplace Safety and Health; in the Manitoba Department of Environment, Mr. Spiegel is Director of Planning. This paper, submitted to the Journal February 26, 1990, was accepted for publication October 19, 1990.

veillance reported from 10 high-risk workplaces. Five workplaces had regular medical monitoring programs throughout the period 1979–87 inclusive, and five conducted comprehensive and continuous programs for at least two years during this period. In other workplaces the medical monitoring was sporadic and voluntary. All blood lead analyses were conducted at the provincial government's Technical Services Laboratory located in Winnipeg, Manitoba. The laboratory is accredited by the American Industrial Hygiene Association and participates in the Blood Lead Proficiency Testing program of the Centers for Disease Control in Atlanta, as well as the program conducted by the Centre de Toxicologie in Quebec. Blood leads are analyzed using the Fernandez Method.[4]

Those whose blood lead concentrations exceeded 2.41 μmol/L (50 μg/dl) on at least one occasion, had an annual frequency of testing ranging from one to 21, median of four and mode of four, whereas those whose blood lead concentrations were always less than 1.93 μmol/L (40 μg/dl) had an annual average testing frequency ranging from one to nine, with median of one and mode of one. To adjust for this variation, a quarterly median was taken for each individual. If an individual had only one blood lead test in a given year, the same result was entered in all four quarters for that year. If an individual had blood tests monthly, only the median of each quarter was entered for analysis.

As measures of blood lead concentrations tend to be skewed, most of the analyses were smoothed (using a median quarterly or annually as the case may be) and the overall medians analyzed. Contingency table analysis allowed the comparison of aggregated distributions of blood lead concentrations, while linear regression analysis assessed company trends over the nine-year period, and Mantel-Haenszel trend tests evaluated specific target proportions over shorter time segments (i.e. proportion of values over 70 μg/dl during the three years following 1979 and proportion over 60 μg/dl in the three years following 1983).

To gain insight into the mechanism underlying the longitudinal variations of individual blood lead concentrations, first order auto correlations to assess the influence of one result on the next among individual workers and testing the run of signs of first differences for randomness to assess individual patterns were performed.[5,6] The overall effect of administrative controls as hypothesized by a negative association between last blood lead concentration nearing the target level and current blood concentrations was tested using the Mantel-Haenszel chi-square stratified by regulatory period.

Industrial hygiene observations and measurements were gleaned from the records of the provincial government's workplace safety and health inspectorate.

Occupational Lead Exposure and Regulatory Process

Workplace safety and health is a provincial responsibility in Canada, with each province enacting its own legislation and regulations. The Workplace Safety and Health Act in Manitoba, proclaimed in 1977, empowers the minister to issue an order requiring an occupational health service at designated workplaces, with penalty of fine for noncompliance. This permitted the minister in charge of occupational health for the province to issue an order to five companies in 1979 to reduce blood lead concentrations to below 3.86 μmol/L, (80 μg/dl). There was no mention of environmental monitoring in the order. An individual was to be removed from the plant when his blood lead concentration exceeded 3.86 μmol/L or transferred to a low lead exposure area when the blood lead ranged from 3.38–3.86 μmol/L (70–80 μg/dl). In 1983, after considerable debate, the 1979 order was revised reducing the blood concentration removal level to 3.35 μmol/L (70 μg/dl), with workers to be transferred to low lead areas when the blood concentration ranged from 2.90–3.38 μmol/L (60–70 μg/dl). There was no mention of differential levels for women, or of environmental monitoring.

The five firms that received this order consisted of one battery manufacturer, two primary lead-using industries, one secondary lead smelter, and one combined battery breaking/secondary smelter operation. These companies all operated continuously from 1979 to 1987 with the exception of the combined battery breaking/secondary smelter operation which was shut down briefly for economic reasons during 1983. Three other companies conducted biological monitoring over part of the same period; a primary copper smelter with additional production of zinc and lead concentrate (1982–87), a battery breaking operation in a reclamation yard (1985–87), and an aluminum forging operation (1985–86) with incidental lead exposure.

In 1986 following years of controversy and no written policy, the Workers' Compensation Board (WCB) in Manitoba essentially established a blood lead concentration of 2.41 μmol/L (50 μg/dl) as a compensable level.[7] Sixty-three claims for lead poisoning had been allowed by the Manitoba WCB prior to 1986; 54 were allowed in the two years following this policy statement.

Results

Table 1 contains the 10,190 blood lead concentrations of employees who worked in the 10 designated plants between 1979 and 1987 inclusive. The proportion of values over 3.86 μmol/L (80 μg/dl) has steadily decreased; there is a statistically significant difference in the distribution of lead concentrations over the years, ($p < .0001$). There were no essential changes in this distribution when one entry per quarter for each worker is included regardless of the number of tests performed on that worker during that particular year (data available from author on request). Figure 1 illustrates the trends in median blood lead concentrations over time in each of the five companies that operated throughout this period (with the exception of a brief plant closure in 1983 in one). Median blood lead levels by year and type of plant are given in the Appendix.

A separate analysis was conducted to assess the proportion of blood lead concentrations over 3.38 μmol/L (70 μg/dl) in the three years following the 1979 order which required a reduction to below this level. Both the overall trend as well as the trend within each of the five firms subject to this order showed a highly significant decrease in the proportion of blood leads over this level (overall from 7.9 percent in 1979 to 0.8 percent in 1982). When a similar assessment was done for the 1983 order, only one of the then six companies showed a statistically significant decrease in proportion of blood leads over 2.90 μmol/L (60 μg/dl), the new target level (data available from author on request). Two companies, a secondary smelter and a combined battery breaking/secondary smelter operation, which had 11.2 percent and 11.8 percent of blood levels over 2.90 μmol/L (60 μg/dl), respectively, at the time of the 1983 order, had 13.9 percent and 16.9 percent over 2.90 μmol/L (60 μg/dl), respectively, three years later.

Industrial hygiene investigation revealed that after the first government intervention, some hygiene and/or engineering controls were implemented in all of the firms in question with each firm adopting a unique strategy depending on its condi-

TABLE 1—Percentage Distributions of 10,190 Blood Lead Concentrations in Manitoba 1979–87 Inclusive

Blood Lead Concentration*	1979	1980	1981	1982	1983	1984	1985	1986	1987
<20	11.5	9.0	8.6	21.2	38.5	36.7	16.4	26.2	37.7
20–29	12.1	15.8	20.4	24.0	22.4	21.8	20.5	18.7	14.7
30–39	13.9	16.8	18.5	19.4	14.2	17.0	19.6	16.4	13.1
40–49	15.4	19.4	17.8	14.2	14.5	14.2	20.7	16.3	15.3
50–59	16.5	18.1	14.5	11.9	6.8	7.2	15.0	15.2	10.5
60–69	12.8	9.6	9.2	6.7	2.7	2.2	5.7	5.0	6.8
70–79	8.4	7.0	7.2	1.9	0.8	0.3	1.1	1.4	1.4
≥80	9.4	4.3	3.8	0.7	0.1	0.6	1.0	0.8	0.4
Total Samples	962	1268	1103	699	973	877	1460	1699	1149

*µg/dL. Conversion factor to SI Units (µmol/L) = .04826

tion. Thereafter, companies seemed to be making greater use of administrative controls, i.e. removing workers to lower lead-in-air areas when blood lead concentrations approached the action level, only to return them to the higher areas after their blood lead concentration had fallen. Industrial hygiene inspections found for example, that while the lead-in-air levels had dropped from 300 µg/M^3 to less than 150 µg/M^3 in a primary lead using company, between 1979 and 1983, such a decrease was not observed in the inspection of a secondary lead smelter where environmental levels continued to exceed 300 µgm/M^3. A second secondary smelter demonstrated a reduction of near 150 µg/M^3 on average but some areas, notably the furnace area, demonstrated no substantial reduction in air-borne lead with concentrations approaching 400 µg/M^3 and personal protection being used as a protective strategy. No ongoing environmental or personal monitoring programs were in place and government monitoring was generally complaint-based. Environmental data are thus incomplete.

To further investigate the apparent failure of the 1983 order to produce a continuing decrease in blood lead concentrations, longitudinal analysis was performed for all workers who ever had values of 2.90 µmol/L (60 µg/dl) or greater. A graph for these workers individually showed a marked "titration" pattern, indicating use of administrative controls. Figure 2 shows two typical examples of this pattern. It demonstrates that as a worker's blood lead concentration approaches the action levels he is removed to a lower lead area, and after his blood lead concentration falls, he is probably returned to a higher lead-in-air concentration since his blood lead level rises. This titration pattern in individual workers was present throughout the period of data collection. Its existence was further demonstrated by the significant (p = 0.0001) finding that if a worker's blood lead was less than 2.41 µmol/L (50 µg/dl) his next blood lead would be higher and vice versa. This indicates that an overall decrease in lead exposure in the plant was not occurring because the lead exposure was being averaged over a pool of employees. Four of the eight companies still in operation in 1985 had 18–50 percent of blood lead concentrations over 2.90 µmol/L (60 µg/dl); one other company brought its percentage down to 1.9 percent, while the remaining three companies had had less than 1 percent even prior to the 1983 order. The workforce turnover rates were higher in the four companies with poor compliance than in the four with lower proportion of concentrations over 2.90 µmol/L (60 µg/dl) (data available from author on request). There was much less evidence of titration in the four with better compliance (data available from author on request).

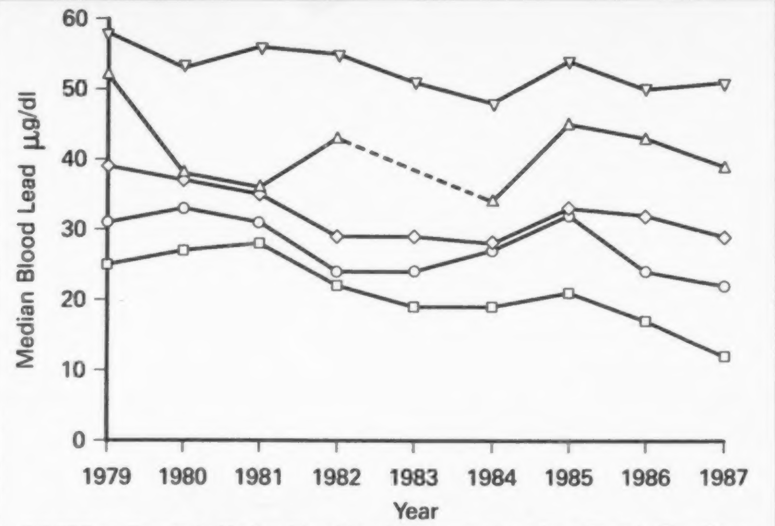

FIGURE 1—Trends in Median Blood Lead Concentrations Over Time in Five Companies Operating during 1979–87

Discussion

The prevention of lead poisoning requires: 1) knowledge of dose response relationships with establishment of an acceptable target blood lead "threshold" to reduce the potential for subclinical lead toxicity, 2) effective surveillance of the environment and of the at-risk population, and 3) action to reduce lead exposure at the source. Considerable knowledge exists with respect to "threshold" levels of exposure for different adverse effects, as summarized elsewhere.[2] There is a need to intensify the monitoring of lead levels in the at-risk population in order to take action which would result in a safer workplace. Methods used to monitor trends of occupational illness and injury[8–13] are usu-

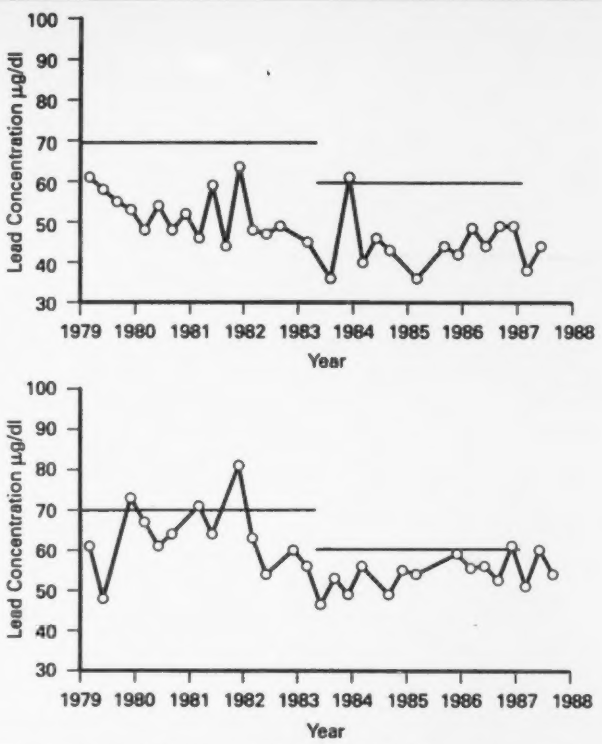

FIGURE 2—Blood Lead Concentrations in Two Individual Workers 1979–87

ally developed primarily for other purposes, making them less useful for occupational health surveillance. Most states have reporting systems specifically for occupational disease,[9] but the completeness of reporting is rarely satisfactory due to a lack of effective centralized monitoring, and disincentives for physicians and others to report occupational disease.[8,9]

One of the main purposes of surveillance systems is to evaluate the effectiveness of worksite control programs.[11] However, there is little information addressing the effectiveness of government regulatory strategy in controlling work-related illness and injury, and there has certainly been controversy[14] in this regard. This study illustrates an overall decrease in the extent of lead poisoning over the last decade. However, on careful examination it is apparent that while the 1979 government order to reduce blood lead concentrations to below 3.38 µmol/L (70 µg/dl) was effective, the 1983 order to reduce blood leads to below 2.90 µmol/L (60 µg/dl) was less so. Thus the conclusion regarding the effectiveness of Manitoba's regulatory strategy is mixed.

Our findings show that when biological monitoring is the only criterion for compliance, utilization of administrative controls enables companies to comply with stricter regulations without decreasing overall lead exposure. The practice of administrative controls is problematic for several reasons. It results in subjecting more people to lead exposure, thereby failing to reduce overall lead exposure; it often causes employment disruption with its consequent socioeconomic consequences; it usually fails to reduce individual body burden of lead. The oscillations observed below the action level (Figure 2) do not represent lead exposure followed by lead excretion. Rather, due to man's relative inability to excrete lead, the downstrokes represent lead storage within the body. The end result of this process is chronic lead poisoning. This finding is supported by the clinical observations of the clinician authors (AY and MT) in which modest blood lead elevations in several long-term employees have been noted to be accompanied by marked increases in erythrocyte protoporphyrin concentrations.

In 1987 the minister responsible for occupational health in Manitoba issued a further order requiring "... a worker to be removed from the lead containing environment until his/her lead level returns to acceptable levels below 2.41 µmol/L (50 µg/dl). The worker may work in a job with minimal lead exposure (30 µg/M^3) considered as an 8 hour average ..." Moreover, the order stated that the exposure to the worker shall be "no greater than 50 µg/M^3 ... Only upon a detailed engineering assessment by a qualified engineer can consideration be given to permit exposure levels ... to exceed 50 µg/M^3." As most Canadian jurisdictions utilize the American Conference of Governmental Industrial Hygienist (ACGIH) level of 150 µg/M^3, this order was considered quite progressive by some and draconian by others. Regardless of the appropriateness of the lead-in-air concentration chosen, the data presented here demonstrate that the inclusion of an environmental target level was a needed addition. Nonetheless, biological monitoring must continue, with the data base described above serving an important role in monitoring the effectiveness of this new regulation.

In conclusion, while appreciation of lead as a subclinical toxin is essential, and a comprehensive blood lead surveillance system for high risk industries is crucial, unless biological monitoring is in conjunction with environment monitoring and measures to decrease exposure at the source, the overall goal of reducing occupational lead poisoning will not be achieved. ☐

Acknowledgments

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APPENDIX—Median Blood Lead vs Year for Lead Exposed Workers from Five Manitoba Companies

	1979–82				1983–87				1979–87
	Observations		P values		Observations		P values		
Firm	Mean	Range	Trend Test[a]	Linear Regression	Mean	Range	Trend Test[b]	Linear Regression	Linear Regression
★ = Secondary smelter	104	80–136	.022	.7706	96	80–108	.364	.1998	.321
△ = Combined battery breaking and secondary smelter[c]	268	128–50	.000	.0001	95	56–148	.618	.0537	.4100
◇ = Battery manufacturer	446	420–452	.005	.0001	526	484–572	.038	.0018	.0265
○ = Primary lead user	174	140–212	N/A	.0001	132	32–168	N/A	.211	.049
□ = Primary lead user	688	504–860	.000	.0001	480	464–504	.471	.1438	.0004

[a] Trend Test for Proportions >70 μg/dl
[b] Trend Test for Proportions >60 μg/dl
[c] Data unavailable for 1983 due to plant closure

Occupational Carpal Tunnel Syndrome in Washington State, 1984–1988

Gary M. Franklin, MD, MPH, Joanna Haug, MSc, Nicholas Heyer, PhD, Harvey Checkoway, PhD, and Neil Peck, MPA

ABSTRACT

Background: There are no published population-based studies of occupational carpal tunnel syndrome (OCTS) using a strict case definition. Most studies are either industry specific or present patient self-report of symptoms.

Methods: We conducted a population-based incidence study of OCTS using the Washington State Workers' Compensation database. Incident OCTS claims were identified with paid bills for physician reported ICD codes 354.0 and 354.1.

Results: There were 7,926 incident OCTS claims identified for the years 1984–1988, which yields an industry-wide incidence rate of 1.74 claims /1,000 FTEs. The mean age (37.4 years) and female/male ratio (1.2:1) in this population differ from those reported in nonoccupational carpal tunnel studies (mean age, 51 years; female/male ratio, 3:1). The female-specific OCTS incidence rate increased significantly during the study period. The highest industry specific OCTS rates were found in the food processing, carpentry, egg production, wood products, and logging industries.

Conclusion: Demographic differences and industry-specific rates consistent with workplace exposures suggest that OCTS is distinct from CTS occurring in nonoccupational settings. Workers' compensation data proved useful in identifying high risk industries. (*Am J Public Health.* 1991;81:741–746)

Introduction

Occupational carpal tunnel syndrome (OCTS) is a major cause of lost work days and workers' compensation costs in the United States. A recent Centers for Disease Control report, based on physician reporting, estimated that 47 percent of carpal tunnel cases are work-related.[1] Industry-specific prevalence rates of OCTS have been reported to be as high as 15 percent.[2] Most such reports reflect prevalence rates in individual plants or small populations of workers.

Population-based data on the incidence of OCTS has been difficult to generate. The principal problem appears to be that OCTS does not have a unique code in the US Department of Labor Z–16.2 coding system. All wrist disorders are aggregated, and specifically diagnosed cases of OCTS cannot be identified routinely. A recent report using Wisconsin workers' compensation data could identify only "potential" OCTS cases.[3]

Washington State Department of Labor and Industries' claims and medical bill payment information systems were utilized to generate age-, sex- and industry-specific OCTS rates for the years 1984–88. In the context of this study, the validity of using such an administrative data base in performing epidemiologic research was investigated.

Methods

Workers' Compensation System

The Washington State Department of Labor and Industries (L&I) is the sole regulator of workers' compensation coverage in Washington State covering approximately 1.3 million full-time workers in 1988. Coverage for roughly one-third of these workers is through self-insurance administered company programs, and the other two-thirds of workers are covered by the Washington State Fund.

Claims Management Data Base

The L&I claims management data base consists of two major data processing systems. The Medical Information and Payment System receives all billing information generated by provider medical bills. This system records such relevant items as dates of service, all associated procedure (CPT) codes, and physician diagnosis by International Classification of Disease (ICD)[4] code for each provider visit.

The Labor and Industries Industrial Insurance System contains all data necessary for administration of claims. Recorded data include age, sex, claim ID, claimant ID, social security number, nature of injury (Z16.2) codes, time loss, and medical aid paid-to-date. In addition, industry type is coded using the Washington Industrial Classification (WIC) system.

From the Occupational Epidemiology and Health Outcomes Program, Department of Environmental Health, School of Public Health and Community Medicine, University of Washington (Franklin, Haug, Heyer, Checkoway); the Department of Medicine (Neurology), School of Medicine, U-WA (Franklin); and the Department of Labor and Industries, Olympia, WA (Peck, Franklin).

Address reprint requests to Gary M. Franklin, MD, MPH, Department of Environmental Health, SC-34, School of Public Health and Community Medicine, University of Washington, Seattle, WA 98195. This paper, submitted to the Journal October 25, 1990, was revised and accepted for publication February 19, 1991.

TABLE I—NIOSH Case Definition for Occupational Carpal Tunnel Syndrome (CTS)

A. One or more of the following *symptoms* affecting at least part of the median nerve distribution: paresthesias, hypoesthesia, pain or numbness; and

B. *Objective findings* consistent with CTS:
Either
  1) *Positive physical examination* findings −(+) Tinel's or Phalen's or decreased or absent sensation to pin prick in median nerve distribution of the hand; or
  2) *Nerve conduction findings* indicative of median nerve dysfunction across the carpal tunnel; and

C. Evidence of work-relatedness—a history of a job involving one or more of the following activities *prior to* development of CTS symptoms:
  1) *Frequent, repetitive* use of the hand or wrist,
  2) Regular tasks requiring the generation of *high force* by the hand,
  3) Regular or sustained tasks requiring *awkward hand positions,*
  4) Regular use of *vibrating hand tools,*
  5) Frequent or prolonged *pressure over the wrist or base of the palm.*

Source = Adapted from Cummings et al.[1]

Incident OCTS Claim Definition

A data extract was generated of all medical bill line items with an ICD code of 354.0 (carpal tunnel syndrome) or 354.1 (median nerve neuritis) from both the claims and medical bill payment systems. The data set included both compensable (time loss) and non-compensable claims for which medical bills were paid between January 1, 1984 and December 31, 1988. Self-insured claims were excluded from both the numerator and denominator in subsequent analyses due to incomplete data.

An incident claim was defined as the first appearance of a paid bill for a claimant with a physician ICD diagnosis of 354.0 or 354.1 between 1984–88. An algorithm was developed to identify unique claimants by claimant ID and social security numbers. This process ensured the removal of multiple claims for the same condition or conditions recurring at a later date, and eliminated problems of missing claimant identifiers for archived claims.

Denominators

OCTS incidence rates by year and industry were estimated directly using worker hours reported to L&I. Worker hours were converted to full time equivalent workers (FTEs) by FTEs = total hours reported/2000.

Age- and sex-specific denominators were not available from worker hours reported to L&I and were estimated by using other available data. Census data for Washington State (1980) were used to estimate an age/sex distribution for the employed population. The Office of Financial Management state and county projected population figures for 1980–91 were used to determine percent changes in the age/sex-specific strata between 1980 and the years of interest.

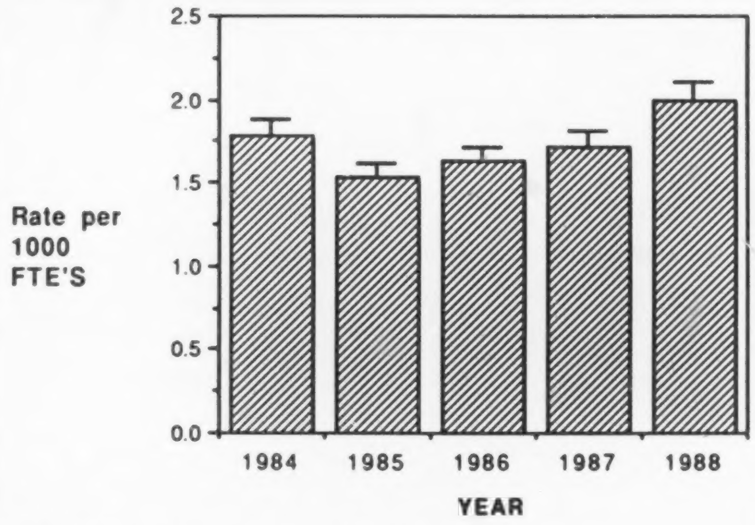

FIGURE 1—Carpal Tunnel Syndrome Incidence Rates 1984–88

Industry-Specific OCTS Rates

Worker hours reported to L&I represent hours of exposure by Washington Industrial Classification (WIC). An industrial classification is a grouping of industries that share similar workplace exposures. Currently there are 975 industrial classifications. OCTS rates are presented using WIC classifications as a surrogate for industry type. In order to eliminate unstable rates, only those classifications with a minimum of 25 OCTS cases over the five-year study period were included in the study.

Misclassification Studies

Two separate sub-studies were performed in an attempt to define potential misclassification: 1) a validation study using the medical records for a random sample of 245 incident claims during the period 1984–88 to estimate (a) the proportion of claims with a documented physician diagnosis, and (b) the proportion of claims that fit the clinical criteria for case definition according to National Institute for Occupational Safety and Health (NIOSH) standards (Table 1)[1], and 2) a review of ICD coding errors associated with a separate sample of L&I claimants who received carpal tunnel surgery between 1984–87.

Statistical Methods

Incidence rates by age, sex, year, and industry are expressed as numbers of OCTS claims/1,000 FTEs. A statistical comparison of rates for male/female differences and test for trend over time were done using formulas reflecting a Poisson distribution.[5,6] Each industry-specific rate was compared to the industry-wide rate, and a crude incident rate ratio with a 95 percent confidence interval was developed.[6]

Results

OCTS Incidence by Age, Sex, and Year

Figure 1 shows that the incidence rate for OCTS claims increased from 1.78/1,000 FTEs in 1984 to 2.00/1,000 FTEs in

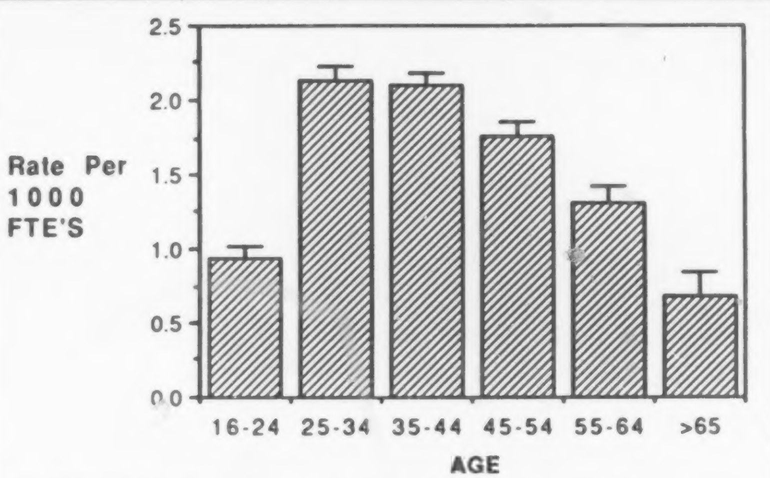

FIGURE 2—Carpal Tunnel Syndrome Incidence Rates by Age 1984–88

1988 (1,533 claims in 1984; 1,320 claims in 1985; 1,475 claims in 1986; 1,642 claims in 1987; and 1,956 claims in 1988). The trend of increase is statistically significant ($p < .01$).

The OCTS incidence rate peaked in the 25–44 age group (mean age 37.4 years) (Figure 2). The overall incidence rate for females (1.96/1,000 FTEs) was significantly greater than that for males (1.58/1,000 FTEs, $p < 0.001$), and this difference has increased over time (Figure 3). The female OCTS incidence rate increased significantly between 1984–88 ($p < 0.005$), while the male incidence rate remained stable with no statistically significant trend.

Industry-Specific OCTS Incidence Rates

For the period 1984–88, the industry-wide OCTS incidence rate was 1.74 claims/1,000 FTEs. The top 20 industries are presented sorted by rate (Table 2) and by total number of incident claims (Table 3). The highest OCTS rate of 25.7 claims/1,000 FTEs for oyster, crab, and clam packing was approximately 15-fold higher than the industry-wide rate (1.74 claims/1,000 FTEs).

Misclassification

OCTS Case-Definition Study—In order to validate the methodology used to identify OCTS claims, medical records for a random sample of 245 incident OCTS claims were reviewed. Two hundred and one (201) out of 245 claims (82 percent) did have a physician diagnosis of OCTS documented in the medical records. The proportion of incident claims without a clear diagnosis of OCTS was consistent across age, sex, and year; no significant differences in proportions were seen within these categories. A high correlation was found between physician diagnosis of OCTS and the criteria set out by NIOSH for case definition. Of the 201 cases with a physician diagnosis in the medical records, 186 (93 percent) met NIOSH standards for OCTS.

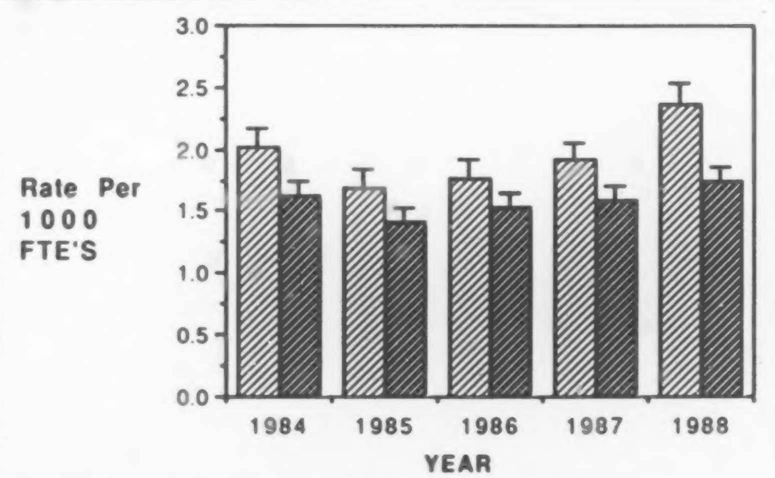

FIGURE 3—Carpal Tunnel Syndrome Incidence Rates by Sex 1984–88

An attempt was made to reduce the potential misclassification due to "false" OCTS claims (missing a clear physician diagnosis) by testing more restrictive incident claim definitions (Table 4). The baseline inclusion criteria of at least one visit with ICD codes 354.0 or 354.1 resulted in a "false" claims rate of 18 percent (44/245) in the validation study. Minimally increasing the inclusion criteria to: a) more than one visit with ICD codes 354.0 or 354.1, or b) subsequent carpal tunnel surgery, would exclude 77 percent (34/44) of the "false" OCTS claims. However, 32 percent (64/201) of the true claims (with physician diagnosis) would also be excluded.

Diagnostic Coding Error Study—In order to determine how much misclassification by diagnosis exists in the data base, an extract was made of all medical bills with a CPT procedure code of 64721 (carpal tunnel decompression surgery) for the time period 1984–87. There were 2,507 unique carpal tunnel surgery claims with 212 associated ICD codes. A classification system (Table 5) was developed to assign the various diagnoses as: 1) CTS with ICD codes 354.0 or 354.1; 2) CTS of traumatic origin; or 3) all other ICD codes. This misclassification study is based on the assumption that all surgeries were indeed cases of CTS. If data were extracted from the system based entirely upon ICD codes 354.0 or 354.1, only 71.6 percent (1,793/2,507) of CTS claims would be captured. Only 3.4 percent (86/2,507) of claims that had received carpal tunnel surgery had diagnosis codes suggesting CTS of traumatic origin. Other diagnosis and data entry errors comprised 25 percent (628/2,421) of CTS surgery claims.

TABLE 2—Rates and Cases of Carpal Tunnel Syndrome by Washington Industrial Classification (WIC), 1984–88 (sorted by rate)

Industrial Class Description	Cases N	Rate/FTE 1000	Rate Ratio	95% Confidence Interval
Oyster, crab, clam packing	50	25.7	14.8	(11.2, 19.5)
Meat, poultry dealers	132	23.9	13.8	(11.6, 16.4)
Packing house	55	18.5	10.6	(8.1, 13.8)
Fish canneries processing	124	18.2	10.5	(8.8, 12.5)
Carpentry	37	11.3	6.5	(4.7, 9.0)
Fruit & vegetable canning	55	10.2	5.8	(4.5, 7.6)
Egg production	30	9.8	5.6	(3.9, 8.1)
Box, shook, pallet bin manufacturing	25	9.3	5.4	(3.6, 8.0)
Sawmills, operation & maintenance	139	9.3	5.3	(4.5, 6.3)
Foundries, steel casting	37	9.0	5.2	(3.8, 7.2)
Logging operations	143	8.9	5.1	(4.4, 6.1)
Wallboard installation	43	8.4	4.8	(3.6, 6.5)
Roofwork, all types	48	8.0	4.6	(3.4, 6.1)
Boat building repair	36	7.7	4.4	(3.2, 6.1)
Plywood manufacturing	54	6.8	3.9	(3.0, 5.1)
Plastic goods manufacturing	121	6.5	3.7	(3.1, 4.5)
Food sundries manufacturing/processing	30	6.4	3.7	(2.6, 5.3)
Wood products manufacturing & assembly	53	6.4	3.7	(2.8, 4.8)
Lumber remanufacturing	26	6.4	3.7	(2.5, 5.4)
Building construction iron/steel/concrete	102	6.3	3.5	(2.9, 4.2)

TABLE 3—Rates and Cases of Carpal Tunnel Syndrome by Washington Industrial Classification (WIC), 1984–88 (sorted by number of cases)

Industrial Class Description	Cases N	Rate/FTE 1000	Rate Ratio	95% Confidence Interval
Convalescent & nursing homes	241	3.1	1.8	(1.6, 2.0)
Grocery/meat stores combination retail	215	3.7	2.1	(1.9, 2.4)
Class H restaurants	207	1.5	0.8	(0.7, 1.0)
Restaurants	170	2.0	1.2	(1.0, 1.4)
Fruit packing	158	5.5	3.1	(2.7, 3.7)
Logging operations	143	8.9	5.1	(4.4, 6.1)
Sawmills, operation & maintanence	139	9.3	5.3	(4.5, 6.3)
Meat, poultry dealers	132	23.9	13.8	(11.6, 16.4)
Fish canneries processing	124	18.2	10.5	(8.8, 12.5)
Plastic goods manufacturing	121	6.5	3.7	(3.1, 4.5)
Building construction iron/steel/concrete	102	6.0	3.5	(2.9, 4.2)
Clerical NOC	92	0.2	0.1	(0.1, 0.1)
Printing, lithography	80	2.9	1.7	(1.3, 2.1)
Fast food, drive-ins	74	1.0	0.6	(0.4, 0.7)
Orchards, fruit	72	1.1	0.6	(0.5, 0.8)
State employees	71	1.8	1.0	(0.8, 1.3)
Clothing manufacturing	71	5.0	2.9	(2.3, 3.7)
Higher education institutions	66	0.4	0.2	(0.2, 0.3)
Electric wiring in buildings	65	1.5	0.8	(0.7, 1.1)
Auto, truck, sales agencies	65	1.4	0.8	(0.6, 1.0)

TABLE 4—Validation Study Using Physician Diagnosis of OCTS

	Physician Diagnosis		
Inclusion Criteria	(Yes)	(No)	
One visit and no surgery	64	34	98
>One visit or subsequent surgery	137	10	147
Total	201	44	245

Discussion

Although findings from a number of recent case-control studies demonstrate a strong association between workplace exposures (high repetition, high force, vibration) and carpal tunnel syndrome,[7–11] this is the first population-based study to report age-, sex- and industry-specific incidence rates of OCTS. The mean age and female/male gender ratio (37.4 years; F:M = 1.2:1) of the OCTS population reported here differ dramatically from that reported for non-occupational carpal tunnel populations (51 years; F:M = 3:1).[12,13] These clear demographic differences between the occupational and non-occupational CTS populations, as well as the magnitude of numbers reported here, would appear to support the concept that workplace exposures may be responsible for a substantial proportion (up to 47 percent)[1] of physician-reported carpal tunnel cases.

The highest incidence rate reported here for shellfish packing (25.7 claims/1,000 FTEs) is similar to the highest reported incidence rates of "potential" carpal tunnel cases using Wisconsin WC data.[3] Industry-specific rates would be expected to vary according to the distributions of industries in each state or region. Some of the highest rates reported here are very characteristic of Washington State (seafood processing and canning, logging and wood products manufacturing, boat building), whereas others are represented in many regions and could be used for comparison (meat and poultry processing, carpentry, fruit and vegetable canning, egg production, foundries, wallboard installation, roofing, plastics manufacturing, and construction).

Industries reported in Table 3, sorted by total numbers of cases, represent more modest incidence rates and rate ratios. Rates for some of these industries (retail groceries, clothing manufacturing, printing and lithography, and electric wiring) may involve relatively weak CTS hazards, or may represent dilution of high exposure jobs by increased job diversification in some industries. Others with very high numbers of claims but non-significant rate ratios (e.g. restaurants) or difficult to explain exposures (e.g. convalescent and nursing homes) require further investigation.

Although video display terminal (VDT) operation has been associated with cumulative trauma disorders,[14] its relationship to OCTS has not been documented clearly in the literature. The industrial classes reported here whose claims may be related to VDT exposure (e.g. clerical, state employees, higher education employees) have non-elevated rate ratios. As these industrial classes represent a wide range of occupations, an association between VDT use and CTS may be diluted. Underreporting of cases in pre-

TABLE 5—Diagnostic Coding Errors for OCTS

71.6%	CTS diagnosis codes 354.0 and 354.1
3.4%	Probable CTS secondary to trauma (fracture, burn)
25.0%	ICD codes not specific for OCTS Miscellaneous data entry errors for OCTS

dominantly female job titles might be an alternative explanation for these lower than expected rates.

Rates among females increased significantly during the study period, while rates among males remained stable. The trends could be due to prior underreporting of claims among females or to a shift of women into higher exposure jobs. Because the gender-specific rates reported here depend on extrapolated rather than direct denominators, these conclusions should be validated in other studies.

Potential pitfalls in the use of workers' compensation[15] and other administrative data bases[16] to generate epidemiologic research have been well documented. Serious underreporting of OCTS cases to workers' compensation bureaus,[1,15] a lack of specificity in the current US Department of Labor coding system,[3] and difficulty generating direct denominators[17] present formidable obstacles that will require public policy changes before improvement is realized. We were able to mitigate the effects of the latter two problems by use of physician-generated ICD billing codes and worker hours reported by WIC, respectively.

Misclassification and coding errors are other serious problems in using workers' compensation data. Incidence rates for OCTS presented were subject to error by both overcounting and undercounting. We were able to validate that only 82 percent of incident claims are true cases of carpal tunnel syndrome by documented physician diagnosis in the medical records. This potential source of overcounting, however, was non-differential by age, sex, or year, and using more or less restrictive inclusion criteria for incident claim definition did not reduce the degree of misclassification.

An equally serious source of potential undercounting of incident claims was also encountered, particularly in regard to ICD coding errors. Only 72 percent of unique claims that were associated with carpal tunnel surgery had an ICD code (354.0 or 354.1) compatible with carpal tunnel syndrome. Another 3.4 percent of claims had diagnosis codes that were related to CTS of traumatic origin, but 25 percent of claims had either non-specific diagnoses or erroneous diagnostic codes. The latter cases would not have been captured in this study. No gains in precision could be derived from applying these estimates of misclassification to the OCTS incidence rates.

The WIC system is specific to Washington State and is not directly comparable to the national Standard Industrial Classification (SIC) coding system. We were unable to generate reliable SIC code-specific OCTS rates because employee hours are not reported to L&I by SIC code. The WIC system differs from the SIC system in that the former sets aside a separate classification for some occupations such as clerical workers regardless of the employers principal industry type. However, the WIC and SIC systems of classification are sufficiently similar to allow other investigators indirect comparability.

On balance, and with more complete understanding of the limitations of workers' compensation[15] and administrative data bases,[16] workers' compensation data can be useful in generating population-based data for the purposes of surveillance, forecasting trends, and resource targeting for loss control and prevention. The use of other surveillance systems, dependent on direct physician or laboratory reporting,[18,19] could provide complementary information. Workers' compensation data might be particularly useful in the investigation of occupational conditions for which no other good surveillance systems exist, such as OCTS, morbid head injuries, labor injuries among minors, sexual assaults,[20] or for conditions related to mandatory reporting, such as lead poisoning.[21] □

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Developmental, Learning, and Emotional Problems Affect 10.7 Million American Children

One in five American children has had a developmental delay, learning disability, or emotional problem according to a survey by the National Center for Health Statistics. The 1988 National Health Interview Survey of Child Health found such disorders among the most prevalent chronic conditions of childhood and adolescence, with a total of 10.7 million children affected. By the teen years, one in four children had experienced one of these disorders and, among male teenagers, almost three in 10 have been affected.

A report on the survey *Developmental, Learning and Emotional Problems: Health of Our Nation's Children, United States, 1988,* states that 4 percent of all children 17 years of age and under (2.5 million) had delays in growth or development; 6.5 percent among children aged 3–17, 3.4 million (6.5 percent) had learning disabilities and 7 million (13.4 percent) have had an emotional or behavioral problem that lasted three months or more or required psychological treatment.

About half of the children with a developmental delay had received treatment or counseling. More than three fourths of those with learning disabilities or emotional problems had received treatment or counseling, according to the parents reporting in the survey. Parents who reported that their child had a learning, developmental or emotional disorder were asked whether the condition made it necessary for the child to attend special classes, or a special school, or get special help at school. About a quarter of children with delays in growth or development and those reported to have emotional or behavior problems received special educational help. About 70 percent of those with learning disabilities received educational assistance for their disabilities.

In relation to family structure, young people from single-parent families or stepfamilies were two to three times more likely to have had emotional or behavior problems than those from two-parent homes. The prevalence of childhood emotional and behavioral problems showed significant variation across family income groups and minority populations with children from less advantaged backgrounds standing a somewhat greater chance of exhibiting such problems. Black and Hispanic parents were less likely than White non-Hispanic parents to report their children's developmental delays or emotional problems, suggesting that emotional and behavioral problems may be underreported by minority parents, when the data from this nationwide household interview survey are compared to data from other sources. Teacher reports and school records, for example, suggest that minority children are over-represented among special education and remedial instruction classes.

The 1988 National Health Interview Survey of Child Health was a survey of 17,100 children 17 years of age and under, residing in the home, and was based on interviews with an informed adult member of the family, usually the mother.

To obtain a copy of the child health report, *Developmental, Learning and Emotional Problems,* contact NCHS, CDC/PHS/DHHS, 6525 Belcrest Road, Hyattsville, MD 20782; (301) 436-8500.

Socioeconomic Status and Survival from Soft-Tissue Sarcomas: A Population-Based Study in Northern Italy

Giovannino Ciccone, MD, Corrado Magnani, MD, Luisa Delsedime, MD, and Paolo Vineis, MD

ABSTRACT

Background and *Purpose:* Differential prognosis among cancer patients according to socioeconomic status (SES) has been reported. We analyzed survival from soft tissue sarcomas (STS) according to different SES indicators.

Methods: We followed up all the adult patients with a new diagnosis of STS occurring between 1.1.1981 and 31.12.1983 in an area of Northern Italy (N = 86).

Results: The overall three-year survival rate was 57 percent. After adjustment for confounders, both low education and blue collar jobs were negatively associated with survival.

Conclusions: The results suggest that patients of low SES have a poorer prognosis for STS. (*Am J Public Health* 1991;81:747–749)

Introduction

The five-year relative survival rates from soft-tissue sarcomas (STS) were about 60 percent in the United States in 1979–84.[1] Several studies have described the clinical and pathological determinants of survival from STS.[2–5]

In recent years, considerable attention has been paid to differential prognosis among cancer patients according to socioeconomic status (SES), particularly for breast and colon cancer.[6–9]

Differences in survival are interpreted as related to earlier access to diagnosis and therapy of higher social class patients.

We are not aware of previous investigations of the role of socioeconomic factors as predictors of survival among STS patients. The present investigation is based on the follow-up of all patients with a new diagnosis of STS among residents of a geographic area of Northwestern Italy.

Methods

A population-based case-control study of STS in the provinces of Novara, Vercelli, and Alessandria in Northwestern Italy was carried out between January 1, 1981 and December 31, 1983 among residents of 20 or more years of age to investigate the role of exposure to phenoxy herbicides in the etiology of STS.[10] Eighty-six subjects were enrolled, after histologic confirmation of the diagnosis by two independent pathologists.[10] The vital status and cause of death of all 86 subjects were ascertained during 1987 through correspondence with the municipalities where they resided at the time of diagnosis. The median follow-up duration was 38 months (mean 34.5; range 0–70).

In 1988 a subset of slides (n = 58) was reviewed by a different group of pathologists to evaluate each case according to a histopathological grading classification proposed by French authors.[3,4]

Within the case-control study, 18 of the 86 original cases could not be interviewed. Details on the questionnaire and the procedures involved are given elsewhere.[10] The following variables were analyzed as potential predictors of survival: age, sex, marital status, education, smoking habits, occupation, exposure to phenoxy herbicides, type of hospital of diagnosis (large public hospital vs small or private hospital), and site of origin, histotype, and histopathological grading of the sarcoma.

The analyses were performed using SAS programs.[11,12] Estimates of survival were obtained using the Kaplan-Meier product-limit method[13]; the statistical significance of the difference between survival curves was evaluated through the log rank test.[14] A multivariate analysis was performed including different SES indicators and other covariates; the Cox proportional hazards model was used.[15] Proportionality of risk over time was assessed for each variable introduced in the models through the plots of the log (−log) of the survival distribution function on the log of time.

From the Unit of Cancer Epidemiology and the Unit of Pathology, Department of Biomedical Sciences and Human Oncology, University of Turin. Address reprint requests to Giovannino Ciccone, MD, Servizio di Epidemiologia dei Tumori, Dipartimento di Scienze Biomediche e Oncologia Umana, Via Santena, 7 10126-Torino, Italy.

TABLE 1—Proportion of Deceased Patients at the end of Follow-up according to Selected Variables Collected from Municipality Rosters or Hospital Records (whole series of 86 cases) and at interview (68 cases)

Variables	Deceased n (%)	Total n
Success of interview		
yes	30(44)	68
no	11(61)	18
Age (years)		
<45	8(36)	22
45–54	7(41)	17
55–64	8(50)	16
65+	18(58)	31
Sex		
males	15(37)	41
females	26(58)	45
Hospital of diagnosis		
large, public	25(50)	50
small or private	16(44)	36
Years of school		
<=5	25(51)	49
6+	4(24)	17
missing	1(50)	2
Marital status		
Married	21(47)	45
Unmarried	4(40)	10
Other	6(46)	13
Smoking		
yes	15(47)	32
no	16(46)	35
missing	0 (0)	1
Main Occupation		
Manual worker	14(52)	27
Farmer	14(58)	24
Others	3(18)	17
Exposure to herbicides		
yes	7(58)	12
no	24(43)	56

TABLE 2—Multivariate Analysis (Cox's proportional hazards model): Role of Education and Occupation in Influencing Survival among the 68 Interviewed Cases*

Model and SES indicators		Coefficient	SE
I	years of school <=5	1.33	0.61
II	manual workers	0.90	0.39
	farmers	0.37	0.43
III	years of school <=5	1.06	0.64
	manual workers	0.65	0.42
	farmers	0.15	0.45

*Each model includes also sex, age, and histopathological grading.

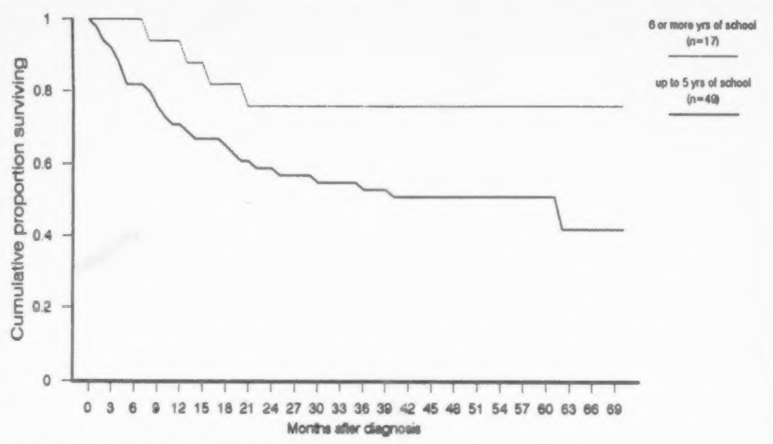

FIGURE 1—Survival from STS by Education

Results

The overall proportion of patients alive three years after a diagnosis of sarcoma of soft tissues was 57 percent.

The Appendix shows the distribution of the patients by site of the tumor, histology, and grading, according to vital status at the end of follow-up. Only two-thirds of cases could be reviewed for histopathologic grading: there is evidence that such classification is likely to be an important prognostic factor for STS.

Table 1 reports the distribution of several characteristics of the patients by vital status at the end of follow-up.

By univariate analysis, the less educated patients and unskilled workers (manual workers or farmers vs others) had poorer prognoses. Figure 1 shows the difference in the proportion surviving according to education. A similar pattern appears comparing survival experiences according to the main occupation.

To take into account the strong association between low education and unskilled jobs and the role played by other covariates in influencing survival, a multivariate analysis was performed, fitting three models. In all models estimates of different SES indicators were adjusted for sex, age, and histopathologic grading. Table 2 shows that low education is negatively related to survival; among blue collar workers, a poorer prognosis is evident for manual workers in industry but not for farmers.

Discussion

Differences in cancer survival according to socioeconomic status of the patient have already been observed for other cancers, both when comparing different categories of workers (i.e. blue vs white collar)[8] and when comparing different ethnic groups.[9] The differences seem to affect particularly breast and colorectal cancers which, at least in theory, could be more influenced by early diagnosis. In a study based on the Swedish Cancer Registry,[8] the following tentative interpretations for the observed social class differences were proposed: earlier detection in higher social class, with improvement of prognosis; earlier detection in higher social class, without improvement of prognosis (lead time bias); host factors, associated with social class, which influence the response to treatment; biological properties of the tumors arising in different social classes.

There is no evidence in favor of host factors or biologic differences in the tumors. Early detection seems to be more reasonable, but it is difficult to disentangle pure "lead time" (i.e. earlier diagnosis without real improvement of survival) and effectiveness of treatment.

Apart from histopathologic grading, only socioeconomic indicators showed an association with survival rates, with patients of low socioeconomic groups showing a worse prognosis. This observation does not rule out a more complex interpretation involving other factors not considered in the present analysis. Although our findings are based on a relatively small number of cases, they cannot be easily explained by bias, since the follow-up was successful for 100 percent of the 86 patients; SES was determined on *a priori* categorization of education and occupation; follow-up procedures were completely blind as to the predictors used in the analyses; differences in survival between patients of different socioeconomic groups were not confounded by age, sex, or tumor grading.

Unfortunately, we had incomplete information about clinical variables, such as stage of cancer at first symptoms or signs or type of treatment. Separate analysis according to histologic type is of little help because there are too few cases in each group. Although they show different probabilities of survival, ranging from zero through almost 100 percent, there is no obvious grouping of histologic types which could make a comparison between social classes easier. The histopathologic grading, a good prognostic factor, was available for only two-thirds of our cases.

The present report, like those cited earlier, suggests the existence of inequalities in health, associated with patient socioeconomic status. More analytical studies to control from a pure "lead time" effect are in order. ☐

APPENDIX—Proportion of Decreased Patients at the End of Follow-up by Site, Histotype,* and Grading of the Soft-tissue Sarcoma.

	Deceased n	Total n
Site of the tumor		
Head and neck	3	8
Retroperitoneal	5	8
Upper limb	4	13
Lower limb	12	23
Thorax	8	13
Abdomen	0	2
Pelvis	4	6
Trunk, unsp.	1	1
Other and unsp.	4	12
Histotype		
Fibrohistiocytic tumors	7	14
Dermatofibrosarcoma protuberans	0	3
Malignant histiocytoma	7	11
Tumors of adipose tissue	2	10
Liposarcoma	0	3
Myxoid liposarcoma	2	7
Tumors of muscle tissue	6	9
Leiomyosarcoma	4	7
Rhabdomyosarcoma	2	2
Tumors of blood vessels	4	17
Angiosarcoma	1	1
Kaposi sarcoma	1	13
Malignant hemangiopericytoma	2	2
Other vascular malignant tumors	0	1
Tumors of synovial tissue	2	3
Synovial sarcoma	2	3
Tumors of peripheral nerves	5	7
Malignant Schwannoma	4	4
Other mal. tumors of peripheral nerves	1	3
Tumors of paraganglionic structures	1	1
Malignant tumors of nonchromaffin tissue	1	1
Tumors of cartilage and bone forming tissue	2	4
Chondrosarcoma of soft parts	2	4
Tumors of uncertain histogenesis	1	2
Malignant cell tumor of soft parts	0	1
Chordoma	1	1
Unclassified	11	19
Histopathologic grading		
Grade I	6	26
Grade II	11	23
Grade III	7	9
Missing	17	28

*For the classification used, see reference 16.

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Public Health Briefs

Pesticide Poisoning Surveillance through Regional Poison Control Centers

Debra Kay Olson, RN, MPH, Lynn Sax, Paul Gunderson, PhD, and Leo Sioris, PharmD

ABSTRACT

The purpose of this study is to describe pesticide exposure in the population of callers to Minnesota Regional Poison Centers. Case files from 1988 reporting pesticide exposure to humans were identified in cooperation with the Minnesota Center for Health Statistics. Data analysis was conducted by computer using SAS statistical package.

Of the 1,428 case files indicating pesticide as the primary substance of exposure to Minnesota residents, a mean age of 5 years (range, one month to 85 years) was identified; 50 percent of all cases were below age 3 years. Males accounted for 1.3 times as many cases as females.

Insecticide was identified in the largest percentage of case files (74 percent) followed by herbicide (12 percent), rodenticide (11 percent) and fungicide-nonmedicinal (3 percent). Ingestion was the most common route of exposure; 85 percent of all calls originated from a residence.

While insecticides are still the most common types of pesticide call, herbicide has surpassed insecticide in production and sales in the US. In this study, herbicide type exposure calls present a much different picture than other pesticide types.

The usefulness of poison control centers for examination of pesticide poisoning is explored. Since reporting occurs coincidental with the exposure and its associated symptoms, each pesticide poisoning report could potentially serve as a true sentinel health event. (*Am J Public Health.* 1991;81:750–753)

Introduction

Pesticides represent a large group of chemicals which in general usage are described in terms of the organism they are intended to kill.[1–3] As the use of these chemicals has unfolded, some disquieting phenomena have been observed. A 1984–85 Nebraska study found an annual incidence rate of pesticide-related illness of 1.35 cases per 10,000 population.[4] Lymphoma and leukemia have been associated with higher death rates among Midwestern farmers in regions of high pesticide use[5,6] and Reeves[7] reported 15 children ages 2 to 17 years as having blood dyscrasias developed shortly after inhalation exposure to household insecticides. Pesticides account for a small but important number of acute human poisonings, particularly in a state with substantial use of pesticides in both rural and urban areas, such as Minnesota.[8,9]

Methods

The purpose of this study was to describe pesticide exposure in the 1988 population of callers to Minnesota Regional Poison Control Centers located at Hennepin County Medical Center and St. Paul-Ramsey Medical Center. Centers document exposure data utilizing the Cooperative Poison Center Report Form as designed by the American Association of Poison Control Centers (AAPCC). Data from these report forms were processed through each Center's computer system and sent to the AAPCC's national data collection project in Denver, Colorado. Data tapes containing all of the 1988 data for each Regional Center were then returned to the Minnesota Center for Health Statistics (MCHS) for collation, formatting, and analysis.

In cooperation with the MCHS all Minnesota resident case files from 1988 reporting pesticide exposure to humans were identified. All calls regarding human contact with a chemical coded by the AAPCC to be a herbicide, insecticide, rodenticide, or fungicide (nonmedicinal) were selected. A total of 2,209 calls were classified by the AAPCC system as pesticide exposure calls for Minnesota in 1988, 4.3 percent of all human exposure calls to Minnesota. Of these calls, 1,428 identified pesticides as the primary substance and 31 call files identified pesticides as a secondary substance of human exposure to Minnesota residents. The remaining 750 calls were incomplete data files or lacked the designation as a Minnesota resident, thus were not available for secondary analysis.

The following descriptive analysis applies to the 1,428 call files indicating pesticide as the primary substance of exposure to Minnesota residents in 1988. These data were derived from self-reported information. The use of poison control caller files imposes the following limitations upon the data: an unknown number of pesticide exposure events may occur for which no call was made; voluntary reporting of data via telephone is verified only by

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Address reprint requests to Debra Kay Olson, RN, MPH, Instructor, Division of Environmental and Occupational Health, School of Public Health, University of Minnesota, Minneapolis, MN 55455. Coauthors: Sax is Research Specialist, and Dr. Gunderson is Director, both at Minnesota Center for Health Statistics; Dr. Sioris is Director, Minnesota Regional Poison Center and Section of Clinical Pharmacy at St. Paul Ramsey Medical Center, and Assistant Professor, College of Pharmacy, University of Minnesota. This paper, submitted to the Journal June 25, 1990, was revised and accepted for publication January 9, 1991.

TABLE 1—Insecticide Exposure Calls by Primary Substance Category, Minnesota, 1988

Substance	Frequency	Percent
Organophosphate only	220	20.8
Organophosphate in combination with carbamate	71	6.7
Organophosphate in combination with chlorinated hydrocarbon	15	1.4
Organophosphate in combination with other pesticide	7	0.7
Chlorinated hydrocarbon only	116	10.9
Chlorinated hydrocarbon in combination with other pesticide	6	0.6
Borates/boric acid (pesticide only)	29	2.7
Carbamate only	189	17.8
Carbamate in combination with other pesticide	6	0.6
Veterinary pesticides (for pets - flea collars, etc.)	49	4.6
Piperonyl butoxide and pyrethrins in combination (without organophosphate or carbamate)	80	7.5
Pyrethrins only (alone)	59	5.6
Rotenone	5	0.5
Piperonyl butoxide only (alone)	2	0.2
Insect repellants (excluding capsicum containing products)	163	15.4
Other	29	2.7
Unknown	14	1.3
Total	1060	100.0

SOURCE: Hennepin Regional Poison Center and Minnesota Regional Poison Center. Secondary analysis performed by the Minnesota Center for Health Statistics and Public Health Nursing, University of Minnesota.

follow-up phone calls by poison center specialists; data acquired by telephone are limited to those who have access to telephones and knowledge of the poison center service; and by AAPCC definition exposure relates only to suspected contact with any substance(s) which, when inhaled, ingested, applied to, injected into, or developed within the body, may cause damage to structure or disturbance of function to living tissue. An actual poisoning event can only be verified through review of medical outcome.

Results

Insecticide was identified in the largest percentage of case files for Minnesota (1,060 cases, 74 percent) followed by herbicide (168 cases, 12 percent), rodenticide (155 cases, 11 percent), and fungicide-nonmedicinal (45 cases, 3 percent). Of the insecticide cases, those containing organophosphates (alone or in combination) were reported most often (Table 1).

Carbamates were reported in 10 cases of fungicide exposure. Of the remaining fungicide cases, 19 were associated with phthalimide, three with non-mercurial fungicides, and 13 unknown/other.

Herbicide cases were most frequently associated with 2,4-D (2,4-dichlorophenoxyacetic acid) or 2,4,5-T (2,4,5-trichlorophenoxyacetic acid) (78 cases, 46 percent) and triazine herbicides (10 cases, 6 percent). Diquat was associated with 13 cases, carbamate with four cases, and other or unknown with 63 cases.

Rodenticide exposure calls were primarily associated with anticoagulant types (128 cases, 83 percent). Strychnine was associated with three rodenticide exposure calls and the remaining 24 rodenticide cases were other/unknown.

Of the 1,428 cases, a mean age of 5 years (range, one month to 85 years) was identified; 50 percent of all calls were regarding children under the age of 3 years (Table 2). The 13–17 years category accounted for the least number of cases overall. The 18+ years old category accounted for 53 percent of all herbicide-related calls.

Male to female ratio was 1.3:1 ± 0.2 for all pesticide-related calls, as well as for individual pesticide types, except herbicides, where males represented twice the number of cases as females.

Ninety-nine percent of all pesticide calls were reported as unintentional poisonings. General calls (not otherwise specified) accounted for 94 percent (1,323 calls) of the unintentional calls while occupational-related exposure made up 4 percent (62 calls).

Calls originating from a residence accounted for 85 percent of all pesticide poisoning calls. Health care facilities accounted for 11 percent and work place for 3 percent.

Ingestion was the most common route of exposure overall, and for each pesticide type (60 percent), except herbicide, where ingestion accounted for 39 percent and dermal exposure accounted for 45 percent of calls.

Pesticide exposure calls by two-hour interval present a bimodal distribution peaking between 10:00 am and 11:00 am (190 calls, 13 percent) and 6:00 pm and 7:00 pm (251 calls, 18 percent). Calls by two-hour interval by pesticide type show a similar bimodal distribution for each specific type of pesticide. The largest percent of pesticide-related calls occur in the months of May (16 percent), July (15 percent), and August (13 percent) (Table 3).

Thirty-three percent of the calls (468 calls) were documented as symptomatic. Two hundred eighty-eight of these were classified as symptomatic related to the pesticide exposure. By pesticide type, rodenticide calls reported the least number (3 percent) of calls as symptomatic related to the exposure.

Ninety-nine percent of all calls were associated with acute exposures. The majority of calls were managed at a non-health care facility (86 percent) while 12 percent were either managed at (107 calls) or referred to (69 calls) a health care facility. More fungicide cases (93 percent) were managed at a non-health care facility (e.g. managed at home) than any other pesticide type.

Discussion

Since 1986, pesticides have accounted for approximately 4.3 percent of all human exposure calls reported to Minnesota Regional Poison Centers. This is consistent with the percent of calls associated with pesticides nationally as reported by the AAPCC.[8] As such, the state data reported here represent but a small portion of a much larger national picture.

Over the past 20 years in the US, organochlorine insecticides have lost favor because of their long-term health effects and persistence in the environment and have been surpassed in use by organophosphates and carbamates.[10] Acute toxicity related to organophosphate pesticides is usually of rapid onset thus symptoms can be more readily associated with acute exposure to this pesticide type. It is thus not unreasonable that exposure to organophosphate insecticides make up the greatest percent of calls (22 percent) in this study. This is also consistent with a Nebraska study where 25 percent of the emergency room visits and hospitaliza-

TABLE 2—Type of Primary Pesticide Exposure Calls by Age, Minnesota, 1988

	Fungicides		Herbicides		Insecticides		Rodenticides	
Age (years)	Freq.	%	Freq.	%	Freq.	%	Freq.	%
<1	4	0.3	0	0.0	98	6.8	16	1.2
1	8	0.5	11	0.7	197	13.8	57	4.0
2	9	0.6	17	1.1	187	13.1	39	2.7
3	4	0.3	7	0.5	91	6.4	15	1.0
4	1	0.1	5	0.4	39	2.7	2	0.1
5	0	0.0	5	0.4	34	2.4	1	0.1
Unknown child								
2–5	0	0.0	5	0.4	19	1.3	1	0.1
6–12	2	0.1	15	1.0	62	4.3	2	0.1
13–17	1	0.1	8	0.5	17	1.2	2	0.1
18+	15	1.0	89	6.3	302	21.1	14	1.0
Unknown age	1	0.1	6	0.5	14	1.0	6	0.5
Total	45	3.1	168	11.8	1060	74.2	155	10.9

SOURCE: Hennepin Regional Poison Center and Minnesota Regional Poison Center. Secondary analysis performed by the Minnesota Center for Health Statistics and Public Health Nursing, University of Minnesota.

TABLE 3—Type of Primary Pesticide Exposure Calls by Month, Minnesota, 1988

Month	Fungicides	Herbicides	Insecticides	Rodenticides	Row	Total (%)
January	1	—	59	7	67	(4.7)
February	1	—	32	11	44	(3.1)
March	2	2	31	11	46	(3.2)
April	7	11	45	18	81	(5.7)
May	13	58	141	17	229	(16.0)
June	7	43	176	12	238	(16.7)
July	5	21	182	6	214	(15.0)
August	4	22	149	14	189	(13.2)
September	3	8	99	19	129	(9.0)
October	1	2	60	15	78	(5.5)
November	—	—	49	12	61	(4.3)
December	1	1	37	13	52	(3.6)
Total	45	116	1060	155	1428	(100.0)

SOURCE: Hennepin Regional Poison Center and Minnesota Regional Poison Center. Secondary analysis performed by the Minnesota Center for Health Statistics and Public Health Nursing, University of Minnesota.

tions for pesticide-related illness during the 1984–85 crop season were due to organophosphates.[4]

In our study herbicide type exposure calls present a much different picture than other pesticide types. They are typically associated with the adult (18+ years) male experiencing a dermal exposure. With the growing use of herbicides both in agricultural (175 percent increase from 25 years ago[2]) and urban settings the percentage of calls received (0.4 percent of all human exposure calls since 1985) in comparison to exposure potential appears low.[11–14]

Looking at the adult environments of work place (six calls) and type of exposure as occupational (18 calls), the difference in exposure for herbicides by age is not explained. It is still general unintentional residence exposures that account for the majority of herbicide calls in this study. Why this overrepresentation of adults for herbicide exposure calls occurs is an area for further research.

With 50 percent of all calls, examined in this study, reporting pesticide exposure to children under the age of three years, concern arises over long-term effects to this population group. Several authors have explored the long-term effects of such exposures on children.[2,15,16]

Chronic exposure to fungicides has been reported by Moses[2] as posing the greatest risk of cancer as compared to other pesticide types. Thus, while representing only 3 percent of the calls in this study, the cumulative effects of such exposures may have far greater implications.

Blanc and Olson[17] suggest that poison centers are ideally suited to occupational disease surveillance as used for early detection of disease, for timely identification of individual cases, and temporal trends. In this sense, each poison center report could potentially serve as a sentinel health event. Sentinel health events refer to a preventable disease, disability, or untimely death whose occurrence serves as a warning signal that the quality of preventive and/or therapeutic medical care may need to be improved.[18,19]

Acute pesticide poisoning reports, as used in this study, would meet the criteria used by Blanc, *et al*,[17,20] in their investigation of the usefulness of poison center data. Those exposures to pesticides with rapid onset of symptoms, such as in the case of organophosphates, are ideally suited for surveillance through such a system. Reporting occurs coincidental with the exposure and its associated symptoms, and there are no delays as in the case of retrospective surveys. In this sense, each pesticide poisoning report could potentially serve as a true sentinel health event for the state. ☐

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Lead Exposure in Outdoor Firearm Instructors

Ram K. Tripathi, PhD, Peter C. Sherertz, PhD, Gerald C. Llewellyn, PhD, and Carl W. Armstrong, MD

ABSTRACT

This study was conducted to determine lead exposure of firearm instructors at an outdoor firing range, while cadets were firing nonjacketed and jacketed lead ammunitions. The breathing zone air for lead exceeded the Occupational Safety and Health Administration standard of 50 μg/m^3 for two instructors during firing exercises using nonjacketed bullets. The use of totally copper-jacketed bullets reduced the breathing zone lead levels by 92 percent for instructor #1 and by 96 percent for instructor #2; subsequent blood lead levels showed a significant decline in both instructors. (*Am J Public Health.* 1991;81:753–755)

Introduction

Lead poisoning from exposure to airborne lead in indoor firing ranges is an occupational disease of public health concern. Several studies of occupational lead toxicity have been documented at indoor firing ranges.[1–7] The major route of absorption for lead in firing ranges is through inhalation of lead dust and fumes. Inadequate ventilation and lack of personal protective equipment use by instructors may result in lead toxicity.

Instructors assigned to firing ranges are likely to have higher exposures because they may spend significant time in that environment. Several occupational studies have suggested that exposure to lead may be a health risk for users of indoor firing ranges.[8–10] However, the extent of this risk among firearm instructors at outdoor firing ranges has not been documented. This study had two objectives: to evaluate the health risks to instructors from airborne lead exposure from nonjacketed, lead bullets; and to document the reduction or elimination of this risk by using totally copper-jacketed lead ammunition.

Methods

Two instructors, not involved in firing, were studied from June 17 to Septem-

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From the Division of Health Hazards Control, Bureau of Toxic Substances, Virginia State Department of Health. Address reprint requests to Ram K. Tripathi, PhD, Virginia Department of Health, Bureau of Toxic Substances, P.O. Box 2448, Richmond, VA 23218. This paper, submitted to the Journal June 11, 1990, was revised and accepted for publication December 5, 1990.

TABLE 1—Personal Breathing Zone Air Lead Levels of Instructors using Nonjacketed and Jacketed Lead Ammunition, 8-Hour, Time-weighted Averages, Richmond, 1987

	Personal Air Lead Level (μg/m³)	
Dates	Instructor #1	Instructor #2
Nonjacketed bullets		
June 18, 1987	36.7	49.1
June 19, 1987	95.6	431.5
July 7, 1987	69.0	152.6
Mean	67.1	211.1
Copper-jacketed bullets		
September 4, 1987	5.4	8.7
Reduction	92.0%	96.0%

TABLE 2—Blood Lead Levels of Instructors using Nonjacketed and Jacketed Lead Ammunition, Richmond, 1987[a]

	Blood Lead Level (μmol/L)	
Dates	Instructor #1	Instructor #2
Nonjacketed bullets		
June 18, 1987	1.01	0.48
June 19, 1987	1.30	0.77
July 7, 1987	1.21	0.77
Mean	1.16	0.68
Copper-jacketed bullets		
September 4, 1987	1.06	0.63

[a] Pre-exposure baseline blood lead levels collected on June 17, 1987, in μmol/L, for instructors 1 and 2 were 1.01 and 0.48, respectively.

ber 4, 1987. Cadets fired a total of 950, 1,539, 3,000 nonjacketed, and 2,160 jacketed bullets on June 18, 19, July 7, and September 4, 1987, respectively. The total number of cadets involved in firing were seven, seven, six, and six on June 18, 19, July 7, and September 4, 1987, respectively. A description of the covered outdoor range was reported previously.[11] Thirty-eight caliber police revolvers and conventional, nonjacketed lead bullets (.38 special caliber, manufactured by 3D Inv, Inc., Doniphan, NE) were used, as well as totally copper-jacketed lead bullets (.38 caliber special ammunition, Omark Industries, Lewiston, IA).

All background, general area, and breathing zone air samples were collected on cellulose ester filters following standard procedures.[11] Background and general area samples were collected within the perimeter of the range one hour before and during firing, respectively. The breathing zone samples were collected near the chest and face of the instructors. All samples were analyzed by atomic absorption spectrophotometry.[12]

Prior to firing activities, individual interviews were conducted and questionnaires administered to obtain information on demographic and other potential sources of lead exposure. Venous blood was drawn to obtain baseline values for blood lead on June 17, 1987. A second questionnaire was administered on July 7, 1987 to assess any symptoms of lead toxicity and to ascertain any other recent potential sources of lead exposure. A third questionnaire was administered on September 4, 1987 after the study was completed. Blood was drawn on June 18, 19, July 7, and September 4, 1987 after each firing activity. All samples were analyzed according to the method of Brodie and Routh (1984)[13] by the National Health Laboratory, Vienna, VA.

Results

The mean concentration of lead in μg/m³ for June 18, 19 and July 7, 1987 was 0.8 (range 0.3–1.2) in the background area air samples and 87 (range 3.8–299) in general area air samples during the firing. The mean concentrations of lead for background and general area air samples during use of copper-jacketed bullets for September 4, 1987 were 0.5 μg/m³ and 9.5 μg/m³, respectively, as previously reported.[14] Use of copper-jacketed ammunition resulted in an 89 percent reduction in lead levels in general area air samples.

The instructors were exposed to different personal breathing zone air lead concentrations on two sampling dates (Table 1). The mean breathing zone lead level with nonjacketed bullets was 67.1 μg/m³ for instructor #1, and 211.1 μg/m³ for instructor #2. The mean breathing zone lead levels during use of the jacketed bullets for instructors 1 and 2 were 5.4 μg/m³ and 8.7 μg/m³, respectively. This resulted in a reduction of 92 percent for instructor #1 and 96 percent for instructor #2.

After use of nonjacketed lead bullets, the mean blood lead level was 1.16 μmol/L for instructor #1, and 0.68 μmol/L for instructor #2 (Table 2). None of these values exceeded the current Occupational Safety and Health Administration (OSHA) return standard of 1.93 μmol/L (40 μg/dl) or removal standard of 2.4 μmol/L (50 μg/dl).[15] After the use of copper-jacketed lead bullets, blood lead levels for instructors 1 and 2 were 1.06 μmol/L and 0.63 μmol/L, respectively.

Discussion

All personal breathing zone lead level samples collected on June 19 and July 7 were above the current OSHA standard of 50 μg/m³.[15] The majority of air lead levels for general area air samples and all breathing zone samples taken while using nonjacketed lead bullets were above the action level of 30 μg/m³, for an eight-hour, time-weighted average (TWA).

The instructor blood lead levels measured during the nonjacketed lead bullet study indicated increases in blood lead concentrations. Blood lead levels reflect very recent lead exposures and are influenced by inhalation and ingestion. They may not reflect the total body burden of lead. Several studies have suggested that inhalation of air containing 1 μg/m³ will increase the blood lead concentrations of an adult by about 0.048 μmol/L when air lead concentrations range from 1 to 5 μg/m³.[16–18] Public health officials, as well as range owners, should be aware of design problems and work habits in outdoor ranges, especially covered ranges where exposures of instructors and other users may occur.

While no definite case of lead toxicity was encountered herein, the instructors experienced increased levels of lead in their blood while cadets were firing nonjacketed lead bullets. Interviews conducted with instructors revealed that the firearm instructors commonly supervise new cadets two to three times per month. This intermittent supervision probably has prevented serious lead toxicity in these instructors. If other covered outdoor firing ranges have air lead levels similar to those found in this study, frequent users would have an increased health risk from the lead exposure based on the present data.

The present study emphasizes a potential health risk of exposure to lead for firearm instructors in this type of outdoor firing range. Such levels were attributed to the firing of conventional, nonjacketed, lead bullets by the cadets. Inadequate ventilation may have contributed to the elevated lead levels found in the general area air samples and personal samples of the instructors. This potential health risk was substantially reduced by using totally copper-jacketed ammunition. The use of jacketed ammunition, as shown in this study and other studies, was very effective in reducing the air lead levels in general area and personal breathing zone air samples.[19–22] Even with such short-time exposures, instructors showed elevated blood lead levels. Remedies to such exposures not only include the use of jacketed ammunition but the adequate ventilation of the facility. □

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Soft Plastic Bread Packaging: Lead Content and Reuse by Families

Clifford Weisel, PhD, Michele Demak, MPH, Steven Marcus, MD, and Bernard D. Goldstein, MD

ABSTRACT

The presence of lead in labels painted on soft plastic bread packaging was evaluated. Lead was detected on the outside of 17 of 18 soft plastic bread bags that were analyzed, with an average of 26 ± 6 mg per bag with lead. Of 106 families questioned, 16 percent of respondents reported turning the bags inside out before reusing for food storage, thus putting food in contact with the lead paint. We estimate that a weak acid, such as vinegar, could readily leach 100 μg of lead from a painted plastic bag within 10 minutes. Further, lead and other metals painted on food packaging of any type becomes part of the municipal waste stream subject to incineration and to landfilling. The use of lead in packaging presents an unnecessary risk to public health. (*Am J Public Health.* 1991; 81:756–758)

Introduction

Low level effects of lead on intelligence and on behavior have been repetitively documented,[1-6] and a causal relation between low levels of lead and fetal malformations, adult hypertension, and auditory acuity has been suggested.[7-10] This recognition has led to efforts to diminish lead exposure, particularly from older homes containing indoor lead paint and from gasoline.

Lead and other metals are also present in colored plastics. However, only submicrogram/cm^2 amounts were extractable in acid or ethanol solutions when these are imbedded in the plastic itself.[11,12] Pigments used to decorate plastics sometimes included lead chromates[13] and if used on food wrappings the potential for transference to food exists.[14,15] While such painting is done on the outside of plastic food wrappings, the opportunity exists for the paint to come in contact with food, particularly if the soft plastic bag is reused. Additionally, the wrappings eventually are included in municipal solid waste disposal facilities and, therefore, the added lead is transferred to the environment.

Methods

Lead Analysis

We purchased 18 breads representing 14 different national brands from three different national chain supermarkets in three New Jersey towns about 80 miles apart: Westfield; Highland Park; and Woodbury. The breads included: white bread, whole wheat, bagels, and pita. A 50 or 200 cm^2 section of each wrapper, containing the printing, was removed, weighed, rinsed with distilled water, and placed in an acid washed beaker. Sections of clear plastic were also cut from several bags to serve as blanks. The total lead content was determined by either a hot or cold concentrated acid extraction. The hot acid extract consisted of 5 mL of nitric acid followed by heating at 250 °C until approximately 0.5 mL remained, then 5 mL HCl with heating until 0.5 mL remain. The cold acid extraction used 5 mL of HCl. The residue from each treatment was diluted to 25 mL with distilled water and analyzed using flame atomic absorption spectroscopy.

Survey Methods

A survey was designed to assess the extent to which soft plastic bread packages may be reused. We targeted families with young children, the subgroup of major concern for increased lead body burden. Two parent groups were surveyed by questionnaire in May and June of 1990. Fifty-three of 84 questionnaires were returned from the first group, and 53 of 90 from the second group (overall response rate 61 percent).

Most respondents are in their thirties, and are well-educated mothers of young children. They are the usual buyers of bread for their households, buying an average of two such packages per week.

Results

Lead was detected in the sections containing printing or designs on all but one of the soft plastic bread bags that were analyzed, but was not present in any of the clear sections of the bags (Table 1). The mean ± S.E. amount of lead for these bags with lead was 26 ± 6 mg. Incubation of the plastic wrappers in 5 percent acetic acid resulted in extraction of 6.8 percent of the lead pigment (mean of 4 studies) in 10 minutes and 10 percent in 60 minutes (mean of 6 studies).

From the University of Medicine and Dentistry of New Jersey-Robert Wood Johnson Medical School. Address reprint requests to Bernard D. Goldstein, MD, Director, Environmental and Occupational Health Sciences Institute, UMDNJ-RWJ Medical School, 675 Hoes Lane, Piscataway, NJ 08854-5635. Dr. Weisel, Ms. Demak, and Dr. Goldstein are also with the Department of Environmental and Community Medicine, and Dr. Marcus is with the Department of Pediatrics, all at UMDNJ. This paper, submitted to the Journal July 20, 1990, was revised and accepted for publication January 8, 1991.

Editor's Note: See also related editorial on page 685 of this issue.

TABLE 1—Amount of Lead in Plastic Food Wrappers from Fourteen Brands of Bread (concentrated acid extraction)

Bread	Amount per painted portion (ug/cm²)	Estimated % of bag with painted writing	Total Lead per bag (mg)
White	7.4	90	13
White	7.8, 27	60	9.4, 32
White	46, 36	80	58, 74
White	14	40	11
White	2, 7.3, 44	95	3.8, 14, 84
White	17	80	27
Pita	17	50	9
Italian	ND	50	ND
Rye	0.9	50	0.9
White	7.9	40	6.3
White	22	95	39
Pita	2.8	50	1.1
White	29	75	44
Rolls	13	40	10

ND—not detected.

An average bag size of 2000 cm² was used to calculate the total lead content on the bags of white bread, 1000 cm² for the smaller bags.

Survey Results

Of the 106 respondents, 41 (39 percent) report reusing the packaging or bag from bread products. The reasons for doing so included: concern for the environment (34 families); convenience (24 families); and saving money (17 families). Food storage and/or packing lunch were the uses most often cited, with more than half of the "reusers" reporting these applications. These packages may be used on multiple occasions: eight families reused them at least three times.

Of the 41 reusers, 23 do so after turning the package inside out, and 17 of these families report using the bags while inside out for food storage. Thus 17 of the 106 respondents (16 percent) store food in contact with the painted label.

Preliminary analyses indicate no relationship between respondents' level of education and package reuse, although respondents who attended graduate or professional school are somewhat more likely to turn the bags inside out. Similarly, no obvious or significant relationship emerges for respondents who are full-time homemakers compared with those working outside the home. Anecdotal information suggests that the rationale for turning the bag inside out before reuse includes removal of bread crumbs that might become moldy, and drying the inside of the bag which picked up moisture from the bread. We were told of paint flaking from the inverted package onto stored food.

Discussion

The Environmental Protection Agency (EPA) estimates that 10 percent of municipal waste is incinerated, 10 percent recycled, and the remainder ends up in landfill.[16] The Food and Drug Administration (FDA) estimates that between 35 and 55 g of white bread and an additional 15 to 25 g of rolls and whole wheat bread are consumed daily across several age groups.[17] Assuming that each package of bread contains one pound and that 60 g of bread are consumed daily, 3.0×10^7 bread bags are disposed daily in the US. Assuming bread bags have an average lead content of 26 mg, 0.8 metric tons of lead would be added to the municipal waste stream daily. This lead could enter the environment as a result of incineration or leaching into groundwater from a landfill. Plastics are currently a prime candidate for recycling. The reduction of toxins is needed[16] to minimize exposure to workers involved in recycling and to increase potential uses of the final product.

EPA estimates that the daily baseline ingestion of lead from food, water, and beverages is between 32 and 40 µg.[17] Inhalation is estimated to give an additional exposure of 20 µg daily.[18] Based on our present findings, we conservatively estimate that a weak acid, such as vinegar, in contact with the pigment for as little as 10 minutes, could extract about 5 percent of the lead. For a 100 cm² surface, the size of a single slice of bread, this would produce an exposure of about 100 µg of lead.

Efforts are underway to reduce lead exposures through the removal of lead from gasoline, solder from cans, and, most importantly, from homes with deteriorated lead paints. However, our findings suggest that an additional source that is dependent upon the habits of individuals, and not measured by current procedures for food analysis, may need to be considered when estimating total lead exposure.

Although presenting an unnecessary risk to health which should be discontinued, we do not believe it likely that the use of inverted bread packaging to store or transport food presents an immediate threat to the health of family members necessitating an emergency response. However, obtaining blood free erythrocyte protoporphyrin or lead levels is a valuable part of the routine medical evaluation of all young children and is particularly important for high-risk groups.[18] Until this use of lead is prohibited, plastic packaging should not be turned inside out to store or transport food. ☐

Acknowledgments

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Non-Hodgkin's Lymphoma in a Cohort of Vietnam Veterans

Thomas R. O'Brien, MD, MPH, Pierre Decouflé, ScD, and Coleen A. Boyle, PhD

ABSTRACT

We examined the incidence of non-Hodgkin's lymphoma (NHL) in a cohort of 18,313 United States Army veterans from the Vietnam era. Diagnoses were confirmed through a review of hospital records. Among veterans who had died after discharge or who had participated in a telephone interview (8,170 Vietnam veterans and 7,564 non-Vietnam veterans), seven Vietnam veterans and one non-Vietnam veteran had developed non-Hodgkin's lymphoma (p = 0.07). As none of the NHL cases had military job titles which suggest that they were occupationally exposed to herbicides while in Vietnam, the reasons for the excess are unclear. (*Am J Public Health.* 1991;81:758–760)

Introduction

Results of studies of occupational groups exposed to phenoxy herbicides have raised concerns that Vietnam veterans may have an increased risk of developing non-Hodgkin's lymphoma (NHL) because of possible exposure to the phenoxy herbicide Agent Orange, used in South Vietnam between 1966 and 1970.[1-3] These concerns have been heightened by reports of an increased risk of NHL among Vietnam veterans.[4,5] In the Centers for Disease Control Vietnam Experience Study (VES), the health of US Army Vietnam veterans was compared with that of Vietnam-era Army veterans who had not served in Vietnam. Although the VES was not designed to assess the risk of rare cancers among Vietnam veterans, the high level of concern about NHL led us to investigate and report the findings for NHL.

Methods

The study population consisted of 18,313 randomly selected, male US Army veterans (9,324 Vietnam veterans and 8,989 non-Vietnam veterans) who entered military service between January 1965 and December 1971. The VES included both mortality and health interview components, and the methods have been described in detail elsewhere.[6,7]

In this analysis, we combined mortality and interview data to determine the number of veterans with NHL in the cohort. Two hundred forty-six Vietnam veterans and 200 non-Vietnam veterans had died between separation from active duty and December 31, 1983, the closing date of the mortality component. A panel of physicians examined medical records of these veterans to determine the causes of death for each veteran.

In the interview component of the VES, we attempted to locate and contact all veterans not identified as deceased. Altogether, we interviewed 15,288 veterans (87.3 percent of eligible Vietnam veterans and 83.8 percent of non-Vietnam veterans) by telephone during 1985–86. If the veteran reported that he had been diagnosed with cancer, we asked him to describe the type of cancer that he had. To verify three self-reported cases of NHL and to identify other cases, we sought medical records for 47 veterans, including men who reported a malignant neoplasm of the lymphatic or hematopoietic tissues, men who reported a diagnosis of cancer of a site that could be NHL (e.g. lung, neck, stomach, lymph nodes), and men who reported a diagnosis of cancer but could not name a specific type or site. We could not

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From the Center for Environmental Health and Injury Control, Centers for Disease Control. Address reprint requests to Thomas R. O'Brien, MD, MPH, CDC, 1600 Clifton Road (E45), Atlanta, GA 30333. This paper, submitted to the Journal May 25, 1990, was revised and accepted for publication January 23, 1991.

TABLE 1—Source, Diagnosis, Age at Diagnosis, Race, Period of Vietnam Service, Military Occupation, Military Region of Service, and Latency Period of Vietnam Veterans Diagnosed with Non-Hodgkin's Lymphoma, Vietnam Experience Study 1985–86

Source	Diagnosis[a]	Age at Diagnosis	Race	Vietnam Service	Military Occupation	Military Region	Months of "Latency"[b]
Interview	Diffuse mixed lymphocytic-histiocytic lymphoma	35	White	1968–69	Infantryman	III	172
Interview	Burkitt's lymphoma[c]	34	White	1968–69	Clerk Typist	III	148
interview	Poorly differentiated lymphocytic lymphoma	25	White	1968–70	Supply	I & III	75
Interview	Malignant lymphoma, large cell type	37	White	1966–67	Radar Maintenance	III	211
Mortality	Burkitt's lymphoma[d]	34	White	1968–69	Artilleryman	II & III	175
Mortality	Lymphoblastic lymphoma	24	White	1968–69	Infantryman	II	47
Mortality	Reticulum cell sarcoma	26	White	1968–69	Administrative Personnel Specialist	III	8

[a] On the basis of medical records.

[b] The "latent period" for NHL was taken as the time period between the beginning of the veteran's tour of duty in Vietnam and the initial medical contact that led to the diagnosis of NHL.

[c] On the basis of review of microscopic tissue specimen, diagnosis was malignant lymphoma (diffuse, small non-cleaved cell type) or Burkitt's lymphoma.

[d] On the basis of review of microscopic tissue specimen, diagnosis was poorly differentiated lymphoid malignancy with extensive sclerosis, consistent with acute lymphoblastic leukemia.

obtain medical records for six of the 47 veterans; none of the six had reported a lymphatic or hematopoietic cancer during the interviews. We made final diagnoses on the basis of pathology reports found in the medical records or, when no pathology report was available, the preponderance of evidence from the medical records. Pathology records were available in 31 of 41 (75.6 percent) of these medical records.

For analytic purposes, we divided the number of veterans with NHL (ascertained through either the mortality component or the interview) in each cohort by the sum of the number of veterans who had died and the number of veterans who had been interviewed (8,170 Vietnam veterans and 7,564 non-Vietnam veterans).

Results

We confirmed that all three veterans who had reported a history of NHL during the interview had been so diagnosed. A fourth veteran who, in the interview, had not specified the type of cancer had also been diagnosed with NHL. All four of these men were Vietnam veterans (Table 1). In the mortality study, three Vietnam veterans and one non-Vietnam veteran were found to have NHL. In total, seven Vietnam veterans and one non-Vietnam veteran were diagnosed with some type of NHL ($p = 0.07$, Fisher exact test (two-tailed)). On the basis of published age-specific cancer incidence data,[8] about 3.9 cases of NHL would have been expected among Vietnam veterans ($p = 0.10$, Poisson) and about 3.5 cases would have been expected among non-Vietnam veterans ($p = 0.14$, Poisson).

All of the Vietnam veterans with NHL had begun their tour in Vietnam from 1966 to 1968. The men had served in a variety of military occupations, and no specific occupation predominated. Regarding pathology, two Vietnam veterans were diagnosed as having Burkitt's lymphoma. On further pathologic review, one case was reported as possibly Burkitt's lymphoma, and the second was reported as acute lymphoblastic leukemia (Table 1).

The "latency period" was short for two Vietnam veterans (Table 1). One veteran first developed NHL eight months after arriving in Vietnam and a second Vietnam veteran had a "latency period" of 47 months.

Discussion

Although these results suggest an increased risk of NHL among Vietnam veterans, our findings should be interpreted with caution. The number of cases is small and, on review, one of the Vietnam veterans may have had lymphocytic leukemia rather than NHL. (These diseases are, however, closely related.[9]) The time period between the beginning of the veteran's tour of duty in Vietnam and the diagnosis of NHL should also be considered when interpreting these data, as the "latent period" for some of the cases presented here is much shorter than the usual period for environmentally caused cancers.[10]

The collective evidence for an unusual risk of NHL among Vietnam veterans is inconsistent. Two other studies have shown that Vietnam veterans may be at increased risk of NHL. Increased mortality from NHL was seen in a large study of US Marine Corps Vietnam veterans (proportionate mortality ratio (PMR) = 2.10; 95% confidence interval = 1.17–3.79), although the same investigators found a deficit of deaths from NHL in US Army veterans (PMR = 0.81; 95% CI = 0.63–1.04).[4] An elevated risk of NHL in Vietnam veterans was also reported in a large population-based case-control study (Odds ratio = 1.47; 95% CI = 1.09–1.97).[5] In contrast, a case-control study using subjects from Veterans Administration hospitals found a modest negative association between Vietnam service and NHL (OR = 0.72; 95% CI = 0.56–0.92).[11]

We could only indirectly assess possible exposure of these veterans to Agent Orange while in Vietnam. Although most of the men served in Vietnam during the time (1967–69) and area (III Corps) of heaviest spraying,[1] three of seven had non-combat military job titles that suggest they were unlikely to have been in contact with herbicides. Although the three men with combat-related job titles may have been exposed to Agent Orange, previous work suggests that, with the exception of men whose military occupation entailed handling herbicides, most US Army combat troops who served in Vietnam were not heavily exposed to dioxin.[12] Since no other factor in the Vietnam experience has been linked to the increased NHL risk, reasons for the excess found in this study and two other studies remain unclear. ☐

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Endemic Giardiasis and Municipal Water Supply

G. Graham Fraser, MB, ChB, MSc, and Kenneth R. Cooke, MB, ChB, PhD

ABSTRACT

To test the hypothesis that endemic giardiasis may be transmitted by unfiltered municipal water supplies, the incidence of laboratory-confirmed giardiasis was studied in a natural experiment due to the arrangement of the public water supply of Dunedin, New Zealand. The incidence rate ratio was 3.3 (90% CI = 1.1, 10.1) for the population receiving unfiltered (microstrained) water relative to that using sand filtered water. In a parallel case-control study of incident cases, the odds ratio for giardiasis and unfiltered (microstrained) water supply was 1.8 (90% CI = 0.5, 6.9). (*Am J Public Health.* 1991; 81:760–762)

Introduction

Epidemics of waterborne giardiasis originating from water supplies have been clearly associated with surface catchments, treatment systems that do not eliminate *Giardia* cysts, and downstream contamination of the reticulation system.[1–5] In contrast, there have been few tests of the hypothesis that surface catchment waters contribute to endemic giardiasis where treatment systems cannot eliminate cysts.[1,6,7]

Giardia cysts have only recently been isolated from water supplies in New Zealand. No substantive epidemics attributable to giardiasis have yet been documented. Many municipal supplies use surface waters and simple treatment methods that would not reliably filter or deactivate giardia cysts.[3,8]

The present study took advantage of a natural experiment—the divided water supply of the city of Dunedin, New Zealand—to test the hypothesis that endemic giardiasis might be transmitted by unfiltered municipal water supplies.

Methods

Dunedin (population 89,000) is supplied almost entirely by water from surface catchments. Most water is filtered by mechanical microstrainers (screen size 23 μm). Part of the city water is treated at a modern station (Mount Grand) using coagulation/flocculation and direct dual media filtration (anthracite and silica sand) that would normally be expected to remove any *Giardia* cysts present.[3,8] All water supplies are chlorinated and fluoridated.

Records of all laboratory-proven cases of giardiasis from persons residing within the Dunedin municipal water sup-

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Address reprint requests to G. Graham Fraser, MB, Manager, Public Health and Primary Care Programs, Otago Area Health Board, Dunedin, New Zealand. He is also Clinical Lecturer, and Dr. Cooke is Senior Lecturer in Epidemiology, in the Department of Preventive and Social Medicine, University of Otago Medical School. This paper, submitted to the Journal June 1, 1990, was revised and accepted for publication January 24, 1991.

TABLE 1—Incidence of Giardiasis in Dunedin: Laboratory-Confirmed Primary Household Cases by Age and Sex*

Age (years)	Cases			Incidence Rate**
	Male	Female	Total	
0–4	5	3	8	7.8
5–14	0	2	2	0.9
15–34	11	15	26	4.3
35–64	5	5	10	2.1
65+	0	1	1	0.5
Total	21	26	47	2.9

*Three cases excluded from this table, as age or sex unspecified.
**Rate per 10,000 person-years.

TABLE 2—Incidence of Laboratory-Confirmed Giardiasis by Water Supply Filtration Method

Filtration Method	Cases*	Incidence**	Relative Risk (90% CI)
Direct dual media granular filtration	2	0.99	1.0 (reference group)
Microstrained	47	3.29	3.3 (1.1, 10.1)

*Primary household cases only. One case could not be classified, as domicile received a blend of both kinds of supply.
**Rate per 10,000 person-years.

TABLE 3—Giardiasis and Exposure to Microstrained Water: Case-Control Study

Exposed to Microstrained Water	Cases	Controls
Ever	19	53
Never	2	10
Odds Ratio (90% CI)	1.8 (0.5, 6.5)	

ply area were obtained from the two medical laboratories in the city for two periods (February to December 1986 and April 1987 to February 1988). Each record included age, sex, residential address, presence of *Giardia* or other pathogens, and the date of receipt of the fecal specimen. The laboratory diagnosis was made by direct microscopy following formalin-ethyl-acetate sedimentation and iodine staining. The denominators for rates were obtained from the 1986 Census of the usually resident populations of the respective water supply areas.[9]

A case-control study was also conducted using cases in the 1987–88 period. Controls were selected from the population referred for laboratory parasitology examination who had negative findings for giardiasis. Three controls were selected per case; cases and controls were matched for age, sex, and season of occurrence of infection. Cases and controls were classified blindly by city council water engineers according to whether their residences were "ever" exposed to microstrained water during the presumed period of infection (defined as seven days to four weeks before the receipt of the fecal specimen).[10]

Relative risks, odds ratios, and confidence limits were computed using the programs of Rothman and Boice.[11]

Results

Fifty primary household cases of giardiasis were diagnosed in the two study periods. Subsequent cases in a household were excluded from the analysis (13 cases). Four further cases were excluded through insufficient or absent information on domicile.

The incidence of laboratory-confirmed giardiasis was greatest in preschool children and in the 15–34 age group (Table 1). The incidence of laboratory-proven giardiasis for residents of the two water supply areas are given in Table 2. In the first study period, 28 of the 29 cases were residents in areas supplied by microstrained water; the remaining case was in an area traditionally supplied by a blend of the two supplies. In the second study period, 19 cases resided in the microstrained areas and two cases in areas that received the Mount Grand supply. The relative risk of laboratory-confirmed giardiasis for residents of areas supplied by microstrained water was estimated at 3.3 (one-tailed P = 0.04). The risk ratio was essentially unchanged following indirect standardization for age.

In the case-control study the odds ratio for exposure to microstrained water was 1.8 (90% CI = 0.5, 6.5) (Table 3).

Discussion

The present findings suggest that endemic giardiasis can originate from the municipal water supply. Both the incidence estimates and the case-control analysis indicate a greater risk of laboratory-diagnosed giardiasis among residents of areas whose supply was treated only by mechanical filtration and chlorination.

The study was small and of correspondingly weak statistical power. As it was limited to the analysis of laboratory reports, information on some possible confounding factors was not available. However, some potential confounders are either unlikely to have had a significant effect, or would have operated as negative confounders. Age differences were not sufficient to account for differences in rates observed. The area with low rates included the most mobile young adult population. The two population areas may have differed in physician practice styles in the investigation of gastrointestinal symptoms, but this should have been accounted for by the case-control study. As most workplaces are supplied by Mount Grand water, misclassification from consumption of cold water at work would be expected to bias the relative risk toward unity. A small proportion of the controls may have had giardiasis (due to the intermittent nature of *Giardia* excretion) but this would bias the odds ratio toward unity.

The consistency of our findings with those of other case-control and descriptive studies[1,6,7] strongly suggests that transmission of endemic giardiasis can occur via municipal water supplies that do not use filtration technology capable of removing *Giardia* cysts. ☐

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The Microbiologic Quality of Drinking Water in North Carolina Migrant Labor Camps

Stephen Ciesielski, PhD, Thomas Handzel, BS, and Mark Sobsey, PhD

ABSTRACT

A two-year study of the microbiological quality of drinking water in 27 randomly selected North Carolina migrant labor camps yielded total and fecal coliform prevalences of 44 percent and 26 percent, respectively in 1988 and similar but higher prevalences in 1989. Preoccupancy testing by county sanitarians had found virtually no total coliform contamination. These findings suggest that a potential source of contamination existed and that current testing protocols which rely on preoccupancy testing may be inadequate. (*Am J Public Health.* 1991;81:762–764)

Introduction

Although migrant farm workers are known to have a high prevalence of enteric infections,[1–7] many of which are water-borne, the potability of drinking water used by migrants has received little attention. This report presents the results of a two-year study of the microbiological quality of drinking water in North Carolina migrant labor camps.

Methods

In 1988, a random sample of 21 migrant camps was selected from a pooled list of camps in eastern North Carolina. Water was also tested in six additional camps, which were part of a related study of the incidence of enteric parasites in migrant farm workers, and 10 control sites (nearby rural businesses and residences). In 1989, all 23 inhabited study camps from the previous year were retested, along with seven additional nonrandomly selected camps. In all camps in both years, water was tested by county health departments prior to occupancy, the results of which were obtained after field study was completed.

Water samples were collected from taps (without aerators) into sterile 1 liter sampling bottles, kept chilled and analyzed within 24 hours. Microbiological testing of 100 ml sample volumes for total and fecal coliforms and *E. coli* was done according to standard protocols using membrane filter methods.[8] Sites with initial positive tests were retested when possible.

Fisher's Exact Test (two-tailed) and Chi-square tests were employed in data analysis.

Results

1988 Results

The bacteriological results for water samples from the 27 camps and 10 control sites are summarized in Table 1; 12 (44 percent) were positive for total coliforms with counts ranging from 1–186 colony forming units (CFU). Eight of these sites were retested and all were positive again. Seven samples (26 percent) were also positive for fecal coliforms. Two of 10 samples were positive for *E. coli*. All 10 control sites were negative for total and fecal coliforms. The presence of latrines (as opposed to flush toilets) was associated with total coliform occurrence in water (p = 0.024; Fisher's Exact Test).

Address reprint requests to Stephen Ciesielski, PhD, Department of Parasitology and Laboratory Practice, School of Public Health, University of North Carolina, Chapel Hill, NC 27599. Mr. Handzel and Dr. Sobsey are with the Department of Environmental Sciences and Engineering at UNCSPH. This paper, submitted to the Journal April 26, 1990, was revised and accepted for publication January 23, 1991.

TABLE 1—Results of 1988 Microbiologic Testing of Drinking Water from Migrant Labor Camps and Control Sites

Sample Source	# Samples	Analyte	# (%) Positive	Mean Bacterial Concentration	
				Geometric	Arithmetic
Migrant Camps	27	T.C.*	12(44)	8.6	16.0
	27	F.C.**	7(26)	0.92	2.65
	10	*E. coli****	2(20)	1.81	4.5
Control Sites	10	T.C.	0	0	0
	10	F.C.	0	0	0

*Total Coliforms.
**Fecal Coliforms.
***Testing for *E. coli* was only performed for 10 of 12 camps with positive coliform results.

TABLE 2—Results of 1989 Microbiologic Testing of Drinking Water from Migrant Labor Camps

Sample Source	# Samples	Analyte	# (%) Positive	Mean Bacterial Concentration	
				Geometric	Arithmetic
Migrant Camps	30	T.C.*	15(50)	12.0	31.93
	30	F.C.**	11(37)	2.19	6.18
	30	*E. coli*	4(13)	0.93	1.0

*Total Coliforms.
**Fecal Coliforms.

1989 Results

Fifteen of 30 camps (50 percent) had water samples positive for total coliforms, 11 (37 percent) for fecal coliforms, and four for *E. coli* (Table 2). Fifty-six percent of resampled camps had the same result as the previous year. In 1989, latrines were not significantly associated with coliforms, but there was an increasing prevalence of coliforms in water samples as a function of time of occupancy. When the results of the four successive samplings of water are grouped as two early samplings and two late samplings, the total coliform prevalence in the early and late groups is significantly different by Chi-square analysis ($p = 0.056$). Rainfall during the week previous to the last two sampling periods totaled 3.68 inches while totaling 0.78 inches for the week prior to the first two sampling periods.[9]

In 1988 only one and in 1989 none of the camps studied were positive for total coliforms on preoccupancy testing by local health departments.

Discussion

In the United States over one-half of the reported outbreaks of waterborne disease are due to fecally contaminated groundwater, most often resulting from overflow or seepage of on-site septic systems.[10] Migrant camps are often overcrowded so that the amount of human fecal wastes may overload the design capacity of on-site sanitary disposal systems. These septic systems may also have exceeded their expected life span. Furthermore, migrants probably excrete greater numbers of pathogens into the septic waste system. Enteric infections, especially parasites, are major health problems among migrants.[1–7] Diarrhea is reported to be 20 times more frequent among migrants than the urban poor.[4] Most intestinal parasites identified in the migrant population can be transmitted by fecally contaminated water[1,3,5,11] and enteric bacteria and viruses have a much greater capacity for waterborne transmission.[7] The concurrent study of the incidence of intestinal parasites in which some study camps were included, found a significant association ($p = .024$) between incidence and total coliforms.* In Florida an echovirus (type 22/23 complex) was isolated from water samples in a migrant camp during an outbreak of gastrointestinal illness.[12]

The surprisingly high prevalence of total and fecal coliforms in migrant labor camp drinking water is a matter of serious concern. According to the US Environmental Protection Agency, "the presence of fecal coliforms in drinking water indicates that an urgent public health problem exists."[14] New standards proposed by the EPA would require the state to be notified within 48 hours when any number of fecal coliforms are demonstrated.[13] The health hazards associated with unsafe drinking water in migrant camps were demonstrated in 1973, when the largest outbreak of typhoid fever in the US since 1939 resulted from the contamination of well water in a Florida migrant camp, preceded by repeatedly high coliform counts in the camp's water supply.[14]

North Carolina's migrant housing regulations require that water in migrant camps housing 13 or more workers be tested before occupancy for total coliforms.[15,16] The failure of pre-occupancy inspections to ensure adequate water quality during habitation probably results from the fact that if coliforms have entered the water system when the camp is occupied in late summer, results are negative the following June (after a nine-month vacancy) due to bacterial die-off and dilution and dispersion in ground water. This variation with usage is supported by the increase in coliform prevalence observed in 1989 with each round of sampling, which suggests that overload of on-site waste water treatment facilities may be a factor. Increasing rainfall in the early autumn may also play a role.

On the basis of these findings, it seems logical to suggest that current regulations be revised to include water quality testing both before and during the period of habitation, not only in North Carolina, but in all states in which migrants are served by ground water sources. Further study is needed to determine the factors responsible for the presence of coliforms in drinking water so that effective intervention can be initiated. ☐

Acknowledgments

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Salmonella Egg Survey in Hawaii: Evidence for Routine Bacterial Surveillance

Myra R. Ching-Lee, MPH, Alan R. Katz, MD, MPH, David M. Sasaki, DVM, MPH, and Henri P. Minette, DrPH

ABSTRACT

One hundred and six dozen eggs, representing 12 brands, were purchased from Oahu supermarkets and cultured for *Salmonella* using standard FDA (Food and Drug Administration) microbiological techniques. Two enrichment incubation temperatures were used to improve culture sensitivity. Ten cartons (9.4 percent) of the 106 dozen samples had shells positive for *Salmonella*. Seven of the 10 were traced to a single egg processor. Inspection of the facility led to the discovery of malfunctioning equipment used in the egg washing and sanitation process. (*Am J Public Health*. 1991;81:764–766)

Introduction

Large outbreaks of human salmonellosis in the United States, related to egg products and cracked shell eggs, led to the passage of the Egg Products Act in 1970. This law required visual inspection of eggs for cracks as well as mandating pasteurization for all bulk egg products.[1,2] Although this legislation led to an initial decrease in egg-related *Salmonella* outbreaks, they have again begun to rise.[2–5]

Contamination of eggs with *Salmonella* can occur in several ways: egg shells may be contaminated by feces[6]; cleaning procedures may be improper[7]; the interior of the egg may be contaminated by the shell being cracked. Contamination may also be internal, prior to shell formation through transovarian transmission.[5,8] In 1989, an Oahu high school student, as part of a science project, cultured two dozen eggs from a supermarket in Honolulu for *Salmonella*. Nine of 24 eggs (37 percent) were culture positive. The high recovery rate prompted the Hawaii State Department of Health to conduct a more definitive prevalence survey for *Salmonella* among eggs for sale in Oahu supermarkets.

Methods

Eggs representing 12 available brands were collected once each week for an eight-week period from June 26 through August 19, 1989. The sampling unit was defined as a carton of 12 large grade A eggs. The carton closest to the aisle was chosen without opening it. Following collection, the eggs were immediately transported to the Hawaii State De-

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From the University of Hawaii School of Public Health, and the Hawaii State Department of Health. Address reprint requests to Myra R. Ching-Lee, MPH, Epidemiological Specialist, Communicable Disease Division, Epidemiology Branch, 1250 Punchbowl St., Rm. 107, Honolulu, HI 96813. Dr. Katz is Assistant Professor of Public Health Sciences at the University; Dr. Sasaki is Veterinary Medical Officer, Communicable Disease Division, Hawaii Department of Health; Dr. Minette is Epidemiology Consultant to the Hawaii State Department of Health. This paper, submitted to the Journal February 27, 1990, was revised and accepted for publication December 20, 1990.

TABLE 1—Culture Results of Egg Shells from Ten Dozen Eggs Positive for Salmonella Using Two Incubation Temperatures

Incubation Temp. 35 Degrees Celsius				Incubation Temp. 41 Degrees Celsius			
Dozen	Egg Brand	*Plating Medium	Serotype	Dozen	Egg Brand	*Plating Medium	Serotype
1	K	XLD	Braenderup	1	K	XLD	Braenderup
1	K	BGA	Braenderup	1	K	BGA	Braenderup
				2	K	XLD	Oranienburg
3	J	BGA	Oranienburg	3	J	BGA	Oranienburg
				3	J	XLD	Oranienburg
				4	H	BGA	Mbandaka
5	K	BGA	Cerro				
5	K	XLD	Cerro				
6	K	BGA	Oranienburg	6	K	BGA	Oranienburg
6	K	XLD	Oranienburg	6	K	XLD	Oranienburg
7	K	BGA	Ohio	7	K	BGA	Ohio
7	K	XLD	Ohio	7	K	XLD	Ohio
				8	H	XLD	Havana
9	K	BGA	Oranienburg & Braenderup	9	K	BGA	Oranienburg
9	K	XLD	Braenderup	9	K	XLD	Braenderup
10	H	BGA	Montevideo	10	H	BGA	Montevideo
10	H	XLD	Montevideo	10	H	XLD	Montevideo & Livingstone

*Media = Plating media used: BGA (Brilliant Green Agar); XLD (Xylose Lysine Desoxycholate).

partment of Agriculture (DOA) laboratory for testing. No eggs were damaged, cracked, or visibly soiled.

Of the 12 brands available for study, three were from US mainland producers and nine were locally produced. Fourteen dozen eggs were potentially available each week from each of the 12 brands, as two brands sold both white and brown eggs. Both white and brown eggs were sampled when available. Three brands on some weeks were not available at the designated supermarket locations. Unsuccessful attempts were made to locate these brands at other store locations. A total of 106 dozen eggs were sampled and cultured.

Isolation and identification of *Salmonella* from eggs were conducted using methodology from the US Food and Drug Administration,[9] with one variation. In addition to the recommended incubation temperature of 35 degrees Celsius (C) for samples in tetrathionate enrichment medium, a second incubation temperature (41 degrees C) was used.[10,11]

In order to avoid cross-contamination of samples, glassware was autoclaved prior to and between each use. Individual sterile pipettes and sterile gloves were used, and the pre-enrichment media were autoclaved.

The eggs from each dozen were separated into two flasks; one containing the shells and the other containing the magma (white and yolks). All samples were incubated for 18–24 hours in a lactose broth, pre-enrichment medium.[9]

Samples from each flask were then transferred to two flasks containing tetrathionate enrichment broth. One set of samples was incubated at 35 degrees C and the other at 41 degrees C.[10,11] After incubation for 24 hours, one loopful from each flask (magma and shell) was streaked onto plates of Xylose Lysine Desoxycholate Agar (XLD) and Brilliant Green Agar (BGA).[9] Isolates were serotyped using Centers for Disease Control methods.[12]

Results

Salmonellae were detected in 10 cartons (9.4 percent) of the 106 dozen eggs sampled. Positive samples were from shells only. *Salmonellae* were detected in samples from only three of the 12 brands examined. The brand with the highest weekly prevalence (6/8) was locally produced; the second highest (3/8) was a mainland brand; while the third highest (1/8) was a local brand.

Use of two enrichment incubation temperatures resulted in improved *Salmonellae* isolation rates. Three of the 10 dozen eggs testing positive came only from the enrichment medium incubated at 41 degrees C. One positive sample came only from the enrichment medium incubated at 35 degrees C. The remaining six dozen positive samples grew at both temperatures. Serotypes *braenderup, oranienburg, mbandaka, cerro, ohio, havana, montevideo,* and *livingstone* were detected (Table 1).

Discussion

An earlier egg survey done in Hawaii[13] as well as surveys conducted in New York,[14] Missouri,[15] British Columbia, [16] India,[17] and Saudi Arabia[18] demonstrated the prevalence of *Salmonella* on the surface of washed, commercially processed egg shells to be 0–2 percent. However, these surveys used only a single enrichment incubation temperature.

The use of two incubation temperatures for the enrichment medium increased the sensitivity for *Salmonella* detection in the present study. *Salmonella* would have gone undetected in three of the 10 dozen eggs had only the standard incubation temperature been used.

The US mainland brand eggs testing positive for *Salmonella* were traced to a USDA supervised plant in California. Upon inspection, no deficiencies were observed in this plant. The two brands of locally produced eggs were traced to one processor and a single farm. As a result of this study, the Hawaii State DOA inspected the processing plant. Equipment used in the egg washing and sanitizing process was found to malfunction. The company voluntarily agreed to cease operations until the deficiencies were corrected. The eggs were diverted to another facility for processing pending repair of the defective equipment. To date, the facility has not resumed operations as an egg processing plant. It is currently being used as a storage facility (personal communication DOA).

Recent reports of egg-associated salmonellosis have incriminated Grade A shell eggs that were considered to have been properly processed.[2–5,8] While transovarian transmission has also been implicated as a possible cause for the resurgence in egg-borne salmonellosis,[5,8] deficiencies in egg-processing procedures may be of more importance.

In Hawaii, the state DOA is the only state agency with regulatory authority over the egg industry. However, the scope of its activity is limited to the structural integrity of the eggs. Currently, there are no regulations that mandate active surveillance and monitoring of eggs for bacterial contamination. Our data indicate that visual inspection of eggs is not enough. Ongoing egg bacterial surveillance with improved laboratory methods, including the use of two enrichment incubation temperatures to increase *Salmonella* culture sensitivity, would allow for early detection of deficiencies in egg handling and processing, and could eliminate *Salmonella* egg contamination as a potential cause of disease in the community. ☐

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Agricultural Machine-Related Deaths

John R. Etherton, MEA, John R. Myers, MSF, Roger C. Jensen, PhD, Julie C. Russell, PhD, and Richard W. Braddee, BA

ABSTRACT

Analysis of 1980-1985 death certificate data for the United States indicated that an average of 369 occupational deaths per year involved agricultural machinery as the external cause of death. Out of all agricultural machine-related deaths, tractors accounted for 69 percent. Over half of these tractor-related deaths were rollovers. There is a need for public health programs to affect greater use of rollover protective structures (ROPS) on farm tractors. (*Am J Public Health.* 1991;81:766–768)

Introduction

Fatalities among farm machine operators present a conspicuous injury target for public health action.[1–3] Farm tractors are known to be particularly deadly,[4–6] but prevention programs appear to be floundering.

As a first step toward the establishment of priorities for fatality prevention programs in agriculture, we decided to examine the National Traumatic Occupational Fatality (NTOF) data base maintained by the National Institute for Occupational Safety and Health (NIOSH).[7,8] Previous analysis of NTOF data for the years 1980-85 had shown that the Agriculture, Forestry, and Fishing industry had a fatality rate of 20.7 per 100,000 workers, a rate 2.6 times higher than the national average for all industries of 7.9 deaths per 100,000 workers.[7]

Methods

Death certificate data in NTOF were reviewed to identify persons killed while working with machines. Machine-related fatalities were identified by codes E919.0 to E919.9 according to ICD-9. Next, the type of machine involved in each machine-related fatality across all United States industries was ascertained by reviewing the description of the cause of death on the death certificate. All agricultural tractor-related fatalities were then categorized and compared by the Standard Industrial Classification (SIC) for the industry in which the incident occurred.[9] Age at death was determined for those fatalities associated with tractors, augers, hay balers, combines, and other agricultural machines.

TABLE 1—Machine-related Occupational Fatalities, All US Industries, 1980–85

Machine Type	Frequency*	Percent of Fatalities
Agricultural machines, E919.0	2216	43.8
Lifting machines, E919.2	1029	20.3
Mining, earth drilling machines, E919.1	467	9.2
Earthmoving, excavating machinery, E919.7	458	9.0
Other specified machinery, E919.8	441	8.7
Unspecified machinery, E919.9	237	4.7
Metalworking machinery, E919.3	88	1.7
Transmission machinery, E919.6	78	1.5
Woodworking machinery, E919.4	47	0.9
Total	5061	99.8**

*Values are based on the 90.6 percent of case records in the NTOF data file that have ICD codes.
**Sum differs from 100 due to rounding.

TABLE 2—Frequency of Agricultural Machine-Related Fatalities (E919.0), All US Industries, 1980–85

Agricultural Machine	Fatalities	Percent
Tractor	1523	69
Auger/elevator	61	3
Combine	53	2
Hay Baler	42	2
Brush hog/mower	40	2
Loader/skidder	37	2
Corn picker	26	1
27 other machines	108	5
Unspecified machine	326	15
Total	2216	101

*Sum differs from 100 due to rounding.

Results

The search of NTOF cases identified 5,061 machine-related fatalities coded E 919.0 to 919.9 (Table 1). Agricultural machines were involved in 2,216 deaths, an annual average of 369 deaths per year. The second ranking machine category on the list is lifting machines (cranes, forklifts, etc.) which averaged 170 fatalities annually, less than half that of agricultural machines.

The seven types of agricultural machines with greatest fatality frequency are listed individually in Table 2. Machines with a frequency less than 1 percent were combined into one category. The final category includes fatalities in which the type of machine was not specified on the death certificate. These data indicate that tractors are involved in the vast majority of agricultural machine-related fatalities. All other agricultural machines are involved in far fewer fatalities when compared to tractors.

Tractor overturns were the single most important event in fatalities involving tractor operators. Of the tractor-related fatalities in the NTOF data set, 52 percent were rollovers. Another 16 percent involved the victim being run over, such as instances in which an operator fell off the tractor and was run over. About 7 percent of the tractor-related fatalities were associated with the power-takeoff.

Farm tractors are used in many industries. For example, using farm tractors to pull cut timber in logging operations is considered to be a manufacturing industry activity. Of farm tractor-related fatalities, 59.3 percent occurred in the agriculture, forestry, and fishing industry, while manufacturing, services, and construction industries each contributed over 5 percent to the national death toll. Victims over 50 years old accounted for 62 percent of farm tractor-related fatalities, and victims over age 60 accounted for 44 percent of farm tractor-related fatalities.

Discussion

The findings reported here are similar to those reported in studies of farm fatalities in Wisconson[5] and Georgia.[6] All of these investigations found that tractor overturns accounted for the largest proportion of farm deaths (Wisconsin 52 percent; Georgia 76 percent; US 52 percent.)

Education, enforcement, and engineering control are often cited as the avenues for prevention of occupational injuries and fatalities. The mandatory education approach to farm tractor safety has not yet received sufficient emphasis. Some training is available through voluntary organizations such as the 4-H and Future Farmers of America. However such training is voluntary. It is unnecessary to pass a test, as is the case with obtaining a license to operate a motor vehicle, unless the tractor driver is age 14–16 and employed by someone not in the family.[10] For all other farm tractor drivers, no license of any kind is required. Furthermore there is no competency-based certification program for tractor operators.

The enforcement approach also has problems. In its annual appropriations bills for the Occupational Safety and Health Administration (OSHA) in the US Department of Labor, Congress has consistently attached a rider which denies OSHA the authority to inspect farms with fewer than 11 employees. This exempts 89 percent of US farms which may choose to comply or not comply with federal safety standards, free of an OSHA inspection. Unlike some consumer products, which may not be sold in the US if they do not meet regulations of the Consumer Product Safety Commission, the sale of farm machinery is not regulated for safety. Because most farm workers are not covered by OSHA enforcement, regulating farm machine safety at the point of sale has been suggested as a promising way to protect farm workers from machine hazards.

The engineering approach involves efforts at preventing rollovers, as well as protecting operators if a rollover does occur by fitting tractors with rollover protective structures (ROPS). Much effort has already gone into the design of ROPS for tractors, and industry-accepted voluntary standards call for the provision of ROPS on new tractors. OSHA has a requirement that employers must have ROPS on new tractors built since October 25, 1976. One of the difficulties encountered in this approach is the fact that farm tractors tend to be used for so many years that replacement of older tractors with newer, safer tractors is proceeding very slowly. There is also a problem with some farmers removing ROPS from new tractors. Considering all these factors, during the next 10 or 20 years it is unlikely that the percentage of tractors with ROPS will increase rapidly unless

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From the Division of Safety Research, National Institute for Occupational Safety and Health, Morgantown, West Virginia. Address reprint requests to John Etherton, DSR/NIOSH, 944 Chestnut Ridge Road, Morgantown, WV 26505. This paper, submitted to the Journal March 19, 1990, was revised and accepted for publication January 23, 1991.

some significant changes take place. Although engineering has provided the technology for preventing most fatalities due to tractor rollovers, it appears that voluntary implementation is very slow.

In Sweden a very significant reduction in tractor rollover fatalities has been brought about by requiring rollover protection on farm tractors.[11] Between 1961 and 1983, Sweden was able to drop its annual fatality rate for tractor overturning from 12 per 100,000 farmers to 1 per 100,000 farmers. This was accomplished through regulations introduced periodically between 1959 and 1981.

It will take cooperative action, including input from farmers, to develop creative prevention programs to solve this difficult problem. Education, enforcement, and engineering will work if they are supported by the common interest of a united agricultural public health community. Achieving this unity is a crucial goal for the 1990s. □

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A Cross-sectional Study of Pulmonary Function in Autobody Repair Workers

David L. Parker, MD, MPH, Kirsten Waller, MD, MPH, Brenda Himrich, MS, Anita Martinez, MS, and Frank Martin, PhD

ABSTRACT

This study evaluated pulmonary function in workers from 39 autobody repair shops. Based on 152 White male workers with known smoking status, the mean percent predicted FEV_1, FVC, and FEV_1/FVC were 93.6, 96.8, and 96.6, respectively. Twenty-three percent of workers had a FEV_1/FVC ratio less than the fifth percentile. Isocyanate levels ranged from nondetectable to 0.06 parts per million (STEL = 0.02 ppm). No shop had an adequate respiratory protection program. We concluded that there was an increase in abnormal pulmonary function in autobody workers. Three recommendations were made to shop owners: functional paint booths should be maintained, respiratory protection programs should be developed, and isocyanate-free paints should be used. (*Am J Public Health.* 1991; 81:768–771)

Introduction

Approximately 700 autobody repair shops employ an estimated 3,500 individuals in Minnesota (Minnesota Department of Labor Relations, Personal Communication). Autobody repair involves hammering dents, removing damaged parts, welding, grinding, abrasive blasting, filling with styrene putty, sanding, and spray painting.[1,2] Workers may be exposed to hazards including isocyanates, metals in paint, solvents, dusts, and noise.[2,3] Despite possible adverse health effects,[4–7] there has been little epidemiologic investigation of the autobody repair industry.[1,8–10] This study examines pulmonary function in autobody repair workers.

Methods

Sampling Protocol

Autobody repair shops listed in the Twin Cities Yellow Pages were randomly selected for study. Shops were included in the study if: 1) they had 20 or fewer employees; 2) at least half of establishment income was from autobody repair; and 3) they were located within the Twin Cities metropolitan area. Shop owners were asked questions regarding the inclusion criteria. The owners of eligible shops were asked to participate.

A questionnaire was given to each shop manager regarding shop age, size, number of cars repaired, ventilation, and repair materials. Sampling for toluene diisocyanate (TDI), total particulates, and solvents (acetone, benzene, methyl ethyl ketone, toluene, xylene, trichloroethylene, and mineral spirits) was done in all

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Address reprint requests to David L. Parker, MD, MPH, Chronic Disease and Environmental Epidemiology, Minnesota Department of Health, 717 Delaware Street SE, Minneapolis, MN 55440. Ms. Martinez is also with the state health department. At the time of the study, Dr. Waller (now with CDC/EIS) and Ms. Himrich (now with the Department of Risk Management, Hennepin County) were students at the School of Public Health, University of Minnesota. Dr. Martin is Associate Professor in Applied Statistics, U-MN. This paper, submitted to the Journal July 17, 1990, was revised and accepted for publication January 23, 1991.

TABLE 1—Characteristics of 39 Autobody Repair Shops

	Mean	Standard Deviation	Range
Number of Workers	3.9	2.6	1–12
Years in Business	18.8	16.0	<1–70
Age of Building	26.4	21.9	<1–90
Repaired each Month	55.4	45.2	9–200

shops. Total isocyanates [TDI and methylene isocyanate (MDI)] were evaluated in 16 shops, lead levels in nine shops, and silica levels in seven shops. Sampling was done in a central work area for three to six hours. Pumps were calibrated before and after each sampling period. Analyses were done according to the National Institute for Occupational Safety and Health (NIOSH) Manual of Analytic Methods.[11]

A self-administered respiratory and work history questionnaire was given to each worker and reviewed with them for completeness. Using American Thoracic Society[14] criteria, pulmonary function was evaluated at the start and end of the work day with a Collins water-sealed spirometer. Measurements consisted of the forced expiratory volume at one second (FEV_1), forced vital capacity (FVC), and the FEV_1/FVC ratio. Spirometric values were adjusted for ambient temperature, barometric pressure, age, sex, and height.

Data Analysis

The shops and workers were characterized using descriptive statistics. Mantel-Haenszel odds ratios, stratified by smoking status (current, past, never) were computed for categorical variables and combined into a single summary odds ratio. Bartholomew's test for trend was used to evaluate the effect of work practices on the proportion of workers with cough or wheeze. Each of the independent variables (e.g. respirator use while sanding) was ordered as: never, sometimes, always.[15]

Percent of predicted FEV_1, FVC, and FEV_1/FVC were calculated for White males using the equations of Crapo, *et al.*[16] An "abnormal" pulmonary function test was one below the fifth percentile for age, height, and sex. The fifth percentile is a value that is approximately 80 percent of predicted for the FEV_1 and FVC, and approximately 90 percent of predicted for the FEV_1/FEC.[17] A one sample binomial test was used to evaluate the prevalence of abnormal pulmonary function tests. Morning and afternoon pulmonary function were compared using paired t-tests. Predictors of pulmonary function were evaluated using stepwise linear regression.[18]

Results

Shop Characteristics

Of the 128 shops contacted, seven (5 percent) were out of business, and 52 (41 percent) were not eligible. Of the 69 remaining shops, 59 were asked to participate of which 39 (66 percent) agreed. Ten shops were not asked to participate because *a priori* study enrollment goals had been met.

Table 1 shows the characteristics of the 39 shops studied. Two shops were one-person operations, and all but one were independently owned and operated. Even though most shops (72 percent) did sandblasting, this was rarely done for more than one hour per week. Ten shops had paint booths in a poor or nonworking condition.

Ambient levels for silica were below the detection limit for the seven shops tested. All solvent levels were below the current American Conference of Governmental Industrial Hygienists (ACGIH) guidelines. Total particulates ranged from 0.4 mg/m^3 to 14 mg/m^3. Particulate levels exceeded 5 mg/m^3 in four shops, and 10 mg/m^3 in one; the current ACGIH guideline for nuisance dusts.[19] Proton-induced x-ray analysis of ambient dusts revealed the presence of small quantities of 20 metals. No ambient lead was detected in the nine shops that were evaluated.

Isocyanate samples were collected from 32 shops. Two samples were lost due to laboratory error. Eight-hour, time-weighted average total isocyanate levels ranged from not detected to 0.06 ppm (mean = 0.005 ± 0.008). The current ACGIH recommended level for TDI and MDI is 0.005 ppm.[20] The highest and lowest isocyanate levels were from the same shop on different days. In three shops, total isocyanate levels were slightly above current ACGIH guidelines. Two to twelve times the current guideline was exceeded in another four shops.

Of 164 workers, 162 (99 percent) completed questionnaires and 159 (97 percent) performed acceptable morning and 149 (91 percent) acceptable afternoon pulmonary function tests. The mean age was 31.6 years (±11.3), 157 (97 percent) were White and 160 (99 percent) males. The mean years worked in the autobody industry was 11.4 (±9.7). Cigarette smoking was common: 67 (41 percent) current, 30 (18 percent) former, 63 (39 percent) never smokers, and two (1 percent) unknown.

Of the 122 workers who spray painted at least one hour per week, 4 percent never used a respirator while painting, 33 percent sometimes did, and 63 percent always used one. For the 96 workers who did sandblasting these percentages of respirator use were 25 percent, 33 percent, and 42 percent, respectively.

Table 2 shows the number of individuals whose pulmonary function was below the fifth percentile and the mean values of the pre-work percent predicted FEV_1, FVC, and FEV_1/FVC stratified by smoking status. Among 60 workers who had never smoked, five (8 percent) had baseline FEV_1 or FVC measurements below the fifth percentile. Fourteen (23 percent) had an FEV_1/FVC below the fifth percentile ($p < 0.001$). No significant change in pulmonary function was seen between the morning and afternoon shifts. In addition, no relationship was seen between shop isocyanate levels and pulmonary function.

By regression analysis, a nonfunctioning paint booth ($\hat{B} = -8.35$), smoking ($\hat{B} = -3.84$), years worked in the autobody industry ($\hat{B} = -0.20$), hours per week sanding ($\hat{B} = 0.21$), doing paperwork ($\hat{B} = 0.20$), and assembly ($\hat{B} = 0.18$) significantly predicted the percent predicted FEV_1/FVC ($R^2 = 0.19$). Decrements in both the percent predicted FEV_1 ($R^2 = 0.20$) and FVC ($R^2 = 0.11$) were significantly associated with years worked in the autobody industry and a nonfunctioning paint booth.

Summary odds ratios showed a positive, non-significant association between the presence of cough and wheezing* and 1) failure to have a functioning paint booth, 2) failure to use a respirator, and 3) spray painting and/or sandblasting. Using

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*Cough was defined as coughing four to six times per day at least four days per week. Wheezing was defined as wheezing on most days or nights.

TABLE 2—Percent Predicted Morning Pulmonary Function and Number of Workers Whose Pulmonary Function Was Less Than the Fifth Percentile by Smoking Status*

Smoking Category (N)	Mean	Standard Deviation	Number with Abnormal Pulmonary Function (%)**
Never Smoker (N = 60)			
FEV_1	93.6	11.4	5 (8)
FVC	96.8	11.4	5 (8)
FEV_1/FVC	96.6	7.0	14 (23)
Former Smoker (N = 30)			
FEV_1	91.6	13.8	6 (20)
FVC	95.6	13.1	3 (10)
FEV_1/FVC	95.8	7.4	8 (27)
Current Smoker (N = 62)			
FEV_1	90.0	15.3	11 (18)
FVC	96.3	13.6	4 (6)
FEV_1/FVC	93.1	10.0	18 (29)

*A total of 159 workers completed at least one acceptable spirometric examination. Of these, five non-White workers and two workers whose smoking status was unknown were excluded from the analysis.

**Abnormal is defined as corresponding to less than the fifth percentile. For the FEV_1 and FVC this is approximately 80 percent of predicted and for the FEV_1/FVC ratio, approximately 90 percent of predicted.

a chi square test for trend, for nonsmokers, the presence of wheezing significantly ($p < 0.05$) increased across the categories of respirator use (always, sometimes, never) while spray painting and for coughing and wheezing while sandblasting. No significant trends were seen for respiratory symptoms and respirator use while sanding.

Discussion

Two-thirds of all shops contacted agreed to participate. Although no information was available on shops that refused to participate, it is unlikely that they were cleaner than the ones tested. Shops were generally dusty and strong solvent odors were common. Improper respiratory protection was often used, when used at all. These findings are similar to those of Tuskes and Key[1] and Rose, *et al.*[8]

Isocyanate levels were consistent with the mild, but significant, decline in pulmonary function and the presence of respiratory symptoms seen.[20,28] Years working in the autobody industry, and working in a shop with a nonfunctioning paint booth, along with smoking, were associated with a decrement in the FEV_1/FVC indicative of obstructive disease.[30,31] Time spent doing tasks not involving painting or priming (i.e. sanding, doing paperwork, and assembly) were positively associated with the FEV_1/FVC. No relationship was found between pulmonary function and ambient isocyanate levels. This was not surprising given the day-to-day variability in the type and quantity of work done in each shop. The failure to find acute pulmonary changes may be due to the migration of sensitized workers who may have had to find other employment or the fact that it was unusual to visit a shop on a Monday. We did not monitor over the course of the week for delayed asthmatic reactions.[26]

The environment in autobody repair shops may be improved in several ways. Shops should maintain functional paint booths for spray painting, and the spray painting of vehicles in open shop areas should be avoided. Shops should have complete respiratory protection programs and use less hazardous, isocyanate-free paints. ☐

Acknowledgments

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Measuring Physical Activity With a Single Question

Kenneth B. Schechtman, PhD, Benico Barzilai, MD, Kathryn Rost, PhD, and Edwin B. Fisher, Jr, PhD

ABSTRACT

Using 1,004 subjects enrolled in a worksite health promotion program, this report evaluated the validity of a single question about participation in regular exercise. Measured at baseline, this one question had a significant age-adjusted association with body mass index ($p < 0.0001$ in women and $p = 0.001$ in men), HDL cholesterol ($p < 0.0001$ in women), and oxygen capacity ($p = 0.0007$ in women and $p = 0.002$ in men). Thus, one self-reported question can provide useful information about who is and who is not participating in regular exercise. The potential validity of a single exercise question is particularly relevant in complex epidemiologic studies where lengthy questionnaires highlight the importance of brief instruments. (*Am J Public Health.* 1991;81:771–773)

Introduction

Through its association with low levels of HDL (high-density lipoprotein) cholesterol, obesity, and other physiologic parameters, physical inactivity has been linked to increased morbidity and mortality from a variety of diseases. These include cardiovascular diseases,[1] diabetes mellitus,[2] osteoporosis,[3] and even breast cancer[4] and colon cancer[5]. In light of these etiologic associations, the measurement of physical activity has become a standard component of epidemiologic research, with at least 30 different measurement techniques having been described.[6] Unfortunately, many of these are complex instruments that are associated with excessive costs and time commitments that make them impractical in large-scale epidemiologic studies.[6,7] Thus, there is great value in establishing the validity of brief, easily obtained measures of physical activity.

It is the purpose of the present report to discuss the utility of a single exercise question that was generated by the St. Louis Working Hearts Program. That question was "Do you currently participate in any regular activity or program (either on your own or in a formal class) designed to improve or maintain your physical fitness?" We will investigate the validity of this question by evaluating its association with three parameters that are known to be associated with physical activity: body mass index (BMI), HDL cholesterol, and oxygen capacity.

Methods

The data discussed herein were generated by Working Hearts, a four-year collaboration between Washington University in St. Louis and a large midwestern communications corporation. All employees at 17 divisions of the corporation were eligible to participate in a multifaceted two-year cardiovascular risk reduction intervention. The intervention included individualized counseling, worksite promotional activities, and 25 awareness seminars and workshops that focused on weight loss, cholesterol reduction, hypertension, exercise, stress reduction, diet, and smoking cessation. Of particular relevance is the fact that the exercise component of the intervention included discussions and workshops, but no regular exercise program. Data on three of the four variables pertinent to this report—exercise behavior, BMI, and HDL cholesterol levels—were sought from all study participants. BMI was de-

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Address reprint requests to Kenneth B. Schechtman, PhD, Division of Biostatistics, Washington University School of Medicine, 660 South Euclid Avenue, Box 8067, St. Louis, MO 63110. Dr. Barzilai is with the Division of Cardiology; Dr. Fisher is with the Center for Health Behavior Research, both at Washington University; Dr. Rost is with the Department of Psychiatry, University of Arkansas for Medical Sciences. This paper, submitted to the Journal August 2, 1990, was revised and accepted for publication January 23, 1991.

TABLE 1—Oxygen Capacity and Baseline Body Mass Index and HDL Cholesterol by Exercise Status at Baseline (Data are expressed as mean ± standard deviation)

	Exercise Status		
Panel A: Women	No (N = 517)	Yes (N = 154)	Age-Adjusted p-value
Oxygen Capacity[1]	28.8 ± 7.5	32.3 ± 7.2	0.0007
Body Mass Index	26.7 ± 6.1	24.2 ± 4.4	0.0001
HDL Cholesterol	57.0 ± 13.4	61.5 ± 14.3	0.0001
Panel B: Men	No (N = 201)	Yes (N = 114)	Age-Adjusted p-value
Oxygen Capacity[1]	33.8 ± 7.5	40.3 ± 8.3	0.002
Body Mass Index	27.6 ± 4.0	26.1 ± 3.2	0.001
HDL Cholesterol	47.2 ± 12.2	48.2 ± 10.2	0.455

[1] Oxygen capacity data were available on 81 women and 33 men involved in vigorous exercise and 155 women and 35 men not involved in vigorous exercise. The unit of measurements is ml/kg/min.

fined as weight (in kg) divided by height (in meters) squared. Oxygen capacity, in ml/kg/min, was measured on a self-selected subset of 304 subjects using a sub-maximal bicycle test. Results, expressed as mean ± standard deviation, are analyzed separately for males and females. Analysis of covariance determined whether there were significant age-adjusted differences in the three relevant parameters between exercising and non-exercising subjects. To analyze changes over time, the 591 participants who provided both baseline and final exercise data were divided into four groups as follows: Group NN responded "no" at both times (N = 281, 47.5 percent); Group YY responded "yes" at both times (N = 130, 22.0 percent); Group NY reported initiating exercise (N = 118, 20.0 percent); and Group YN reported cessation of regular exercise (N = 62, 10.5 percent).

Results

The Working Hearts data set contained 1,004 subjects whose age was 36.6 ± 9.5 years; 73.1 percent were female. Among the 986 subjects who provided baseline exercise information, 32 percent (N = 316) reported participation in regular exercise. Subjects who provided oxygen capacity data (N = 304) were younger (p = 0.001 in women, p = 0.012 in men), had a lower baseline BMI (p = 0.036 in women, p = 0.049 in men), and were more likely to exercise at baseline among women (p = 0.008) but not among men (p = 0.265), than those not providing oxygen capacity data. However, after age adjustment, logistic regression found that BMI had no independent association with the availability of bicycle test data in either sex, but that baseline exercise was still significant in women (p = 0.030). The availability of bicycle test data was not associated with baseline HDL cholesterol in either sex.

Table 1 provides the key data summarizing the validity of the simplified exercise measure we have utilized. It indicates that at baseline, and after age adjustment, subjects who said they were exercisers had a lower BMI (p < 0.0001 in women and p = 0.001 in men), greater oxygen capacity (p = 0.0007 in women and p = 0.002 in men), and had a higher HDL cholesterol (p < 0.0001 in women) than those who did not report participation in regular exercise.

Using the four groups defined in the Methods section, Table 2 contains the mean change from baseline to final measurement in both BMI and HDL cholesterol. Oxygen capacity was not evaluated in this context because it was measured at only one time point. Following age adjustment, mean HDL cholesterol was not the same in all four groups of men (p = 0.007). Statistical contrasts show that this reflected an *increase* in HDL in group YY (2.2 ± 11.5, N = 58) that was significantly

TABLE 2—Changes in Body Mass Index and HDL Cholesterol by Exercise Status[a] (Data are expressed as mean ± standard deviation)

	Group NN	Group NY	Group YN	Group YY	
Panel A: Women	(N = 209)	(N = 92)	(N = 43)	(N = 69)	Age-Adjusted p-value[b]
Change in body Mass Index	0.59 ± 1.9	0.05 ± 2.2	0.69 ± 2.2	0.23 ± 1.7	0.100
	Contrasts:[c] Group NN vs Group NY: p = 0.028				
Change in HDL Cholesterol	−1.3 ± 10.7	−2.2 ± 6.7	−3.3 ± 11.5	−2.8 ± 7.9	0.655
	Group NN	Group NY	Group YN	Group YY	
Panel B: Men	(N = 72)	(N = 26)	(N = 19)	(N = 61)	Age-Adjusted p-value[b]
Change in body Mass Index	0.33 ± 1.3	−0.47 ± 1.5	−0.02 ± 0.87	−0.08 ± 2.5	0.218
Change in HDL Cholesterol	−3.0 ± 6.4	−1.7 ± 6.4	−1.0 ± 4.9	2.2 ± 11.5	0.007
	Contrasts:[c] Group NN vs Group YY: p = 0.0006				

[a] Exercise groups are: Group NN = No exercise at baseline or final screening; Group NY = Not exercising at baseline, exercising at final screening; Group YN = Exercising at baseline but not at final screening; Group YY = Exercising at both time points.

[b] Based on analysis of variance testing equality in all four groups and adjusting for age.

[c] Not adjusted for multiple comparisons.

different from the *decrease* in HDL cholesterol in group NN (−3.0 ± 6.4, $p = 0.0006$, $N = 70$). Although the overall age-adjusted ANOVA showed no difference in BMI among the four groups of women ($p = 0.100$), it is worth noting that the statistical contrast comparing the increase in BMI in group NN (0.59 ± 1.9) with the small change in group NY (0.05 ± 2.2) was significant ($p = 0.028$).

Discussion

These results indicate that among Working Hearts subjects, a single self-reported measure of participation in regular exercise at baseline was significantly associated, after age adjustment, with body mass index and oxygen capacity in both sexes and with HDL cholesterol in women. Thus, our data support several other studies[8-10] which have suggested that simple self-reported instruments can yield valid information about participation in regular exercise. Based on only two exercise questions, for example, the Lipid Research Clinics Study[8] found that subjects who reported regular vigorous exercise had higher HDL cholesterol, lower resting heart rates, and longer exercise test durations than those not reporting regular exercise. Siconolfi et al[9] found that a single question concerning the self-reported frequency of exercise that induced sweating was associated with maximal oxygen uptake.

If a prime focus of a study is to gain detailed and precise information about levels of physical activity, extensive exercise questionnaires are required. Our results suggest, despite some positive associations in Table 2, that simplified exercise measures may also be inadequate when changes over time are of primary interest. However, as with the reports noted above,[8-10] our data argue strongly that one or two self-administered questions can provide a valid measure of who is and who is not currently involved in regular exercise. Thus, when extensive questionnaires place great demands on participants in a multifaceted epidemiologic study, and when detailed information about exercise behavior is not essential, a few simple exercise questions may be an appropriate alternative to the complex instruments that are frequently utilized. ☐

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Public Health and the Law

OSHA's Four Inconsistent Carcinogen Policies

Background

No issue has proven more contentious in occupational health policy than the control of chemical carcinogens. The Occupational Safety and Health Administration (OSHA) has focused much of its efforts on cancer risks, both through substance-specific regulations and through broader efforts to establish generic policies and work practices. The 1980 Generic Carcinogen Policy was conceptualized as the centerpiece of OSHA's shift from substance-specific to generic regulations covering hundreds if not thousands of substances. Despite adverse judicial rulings and a generally anti-regulatory political climate, OSHA continued to pursue a generic approach to occupational health hazards during the 1980s, particularly through its 1983 Hazard Communication Standard and 1989 Air Contaminants Standard. The legal and political wrangling have left deep scars on OSHA's strategy, however. While consistency and comprehensiveness were once the principles underlying OSHA's efforts to control carcinogens, the current collection of policies and regulations is remarkable for the inconsistent and incomplete way in which suspect chemicals are treated.

In this paper, we analyze OSHA's four carcinogen strategies, as embodied in the Generic Carcinogen Policy, the substance-specific carcinogen regulations, the Hazard Communication Standard, and the Air Contaminants Standard. Two issues are of particular interest.

- Given the inherent methodological limitations to epidemiological data on occupational and environmental cancer, a major scientific and policy debate has developed over the appropriate use in human risk assessment of laboratory evidence on chemical carcinogenesis in animals. The four OSHA policies exhibit quite different positions on this debate.
- The choice between direct regulation and indirect labor market pressure, and the stringency of those exposure limits that are imposed, reflect different attitudes toward the appropriate balance of economic costs and health benefits for occupational health regulations. OSHA's four policies reflect dramatically different balancing efforts.

We begin with the Generic Carcinogen Policy, which remains OSHA's formal policy on the issue but which has been blocked by administrative stays and never incorporated into risk management programs. The substance-specific cancer standards are then analyzed in terms of their comprehensive treatment of a few public health hazards and neglect of many others. While partially an attempt to preempt stronger state and local regulations, the Hazard Communication Standard is in many ways both comprehensive and innovative, covering a wide range of chemicals and imposing duties on employers that could potentially lead to substitution of alternative products and processes. Given its current prominence as the centerpiece of OSHA's regulatory program, the Air Contaminants Standard receives the greatest emphasis. Containing its own internal inconsistencies, the Air Contaminants Standard embodies the best and the worst of OSHA's new regulatory philosophy and its strategy for controlling chemical carcinogens in the workplace.

The Generic Carcinogen Policy

OSHA's Generic Carcinogen Policy[1] was a response to two complementary concerns that emerged over the course of the 1970s and came to dominate the agency's agenda. OSHA had devoted a substantial portion of its resources to the regulation of individual carcinogens but had achieved only meagre results. The agency leadership faced growing demand for an accelerated regulatory timetable from Congress, organized labor, and environmental groups. Of equal significance, perhaps, was the competition among several federal agencies for leadership in formulating governmental policies with respect to chemical carcinogens. Proposed in 1977 and promulgated in 1980, OSHA's Generic Carcinogen Policy embodied a highly protective approach to chemical carcinogenesis. Based on the most conservative set of assumptions concerning methods of cancer induction and the most stringent set of requirements for exposure controls, the 1980 policy represents the high water mark of governmental enthusiasm for regulating occupational and environmental carcinogens.

OSHA devoted much of its energies during the 1970s to developing standards regulating exposure to occupational carcinogens, including asbestos[2] in 1972, vinyl chloride[3] and a group of 14 chemicals[4] in 1974, coke oven emissions[5] in 1976, plus benzene,[6] dibromochloropropane (DBCP),[7] inorganic arsenic,[8] and acry-

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From the Center for Occupational and Environmental Health, University of California, Berkeley. Address reprint requests to James C. Robinson, PhD, Center for Occupational and Environmental Health, School of Public Health, University of California, Berkeley, CA 94720. This paper, submitted to the Journal June 20, 1990, was revised and accepted for publication January 30, 1991.

George J. Annas, JD, MPH, Chair, Health Law Department, Boston University, Schools of Public Health and Medicine, 80 East Concord Street, Boston, MA 02118, is editor of the Public Health and the Law column.

lonitrile[9] in 1978. All of these except the DBCP and acrylonitrile standards underwent court challenges from the regulated industries, with lengthy delays and threats of reversal. A number of "science policy" issues reappeared at each challenge, including the relevance of animal data for human carcinogenesis, the importance of negative study results compared to positive results, the correlation between induction of benign and malignant tumors, and the purported existence of a threshold for carcinogenic effects. These same policy issues were being raised at EPA with respect to pesticides covered by the Federal Insecticide, Fungicide, and Rodenticide Act[10] and toxic air pollutants covered by the Clean Air Act.[11] After some discussions with EPA over the possibility of a joint carcinogen policy, OSHA decided to proceed with its own standard. The history of the OSHA policy and its relationship with initiatives at other regulatory agencies has been described in detail in several publications.[12–17]

The Generic Carcinogen Policy established a categorization system for known or suspected chemical carcinogens, with immediate regulatory action being proposed for substances falling into Category I. This category would be composed of substances that increased the incidence of benign or malignant tumors or decreased the latency period between exposure and onset in humans, in two animal species, in a single mammalian species in two or more independent experiments, or in a single experiment if supported by evidence from short-term tests. Within the universe of substances falling into Category I, the policy established a priority setting mechanism whereby the agency would select 10 substances for comprehensive standard setting at any one time, to include requirements for exposure limits, monitoring, worker training, and other factors. In 1980, OSHA issued a candidate list of 204 chemicals for review as possible candidates under the terms of the policy. The exposure limits were to be set at the lowest "feasible" level, which was interpreted by the agency as accepting any level of compliance costs short of causing massive economic dislocation and job loss in the regulated industry.

The weakening of the Generic Carcinogen Policy began almost as soon as it was promulgated. In July of 1980, the United States Supreme Court upheld the decision by the Fifth Circuit Court of Appeals to vacate the standard governing occupational exposure to benzene, on the grounds that OSHA had not proven that exposures to benzene at the level regulated by the standard posed a "significant risk" to employees.[18] This ruling implied that any standard not demonstrating significant risk, presumably through the emerging quantitative risk assessment methodologies, would be overturned. The types of evidence sufficient to qualify chemicals for OSHA's Category I did not demonstrate "significant risk" without further analysis. In January of 1981, OSHA issued a revised Generic Carcinogen Policy that mandated a two-part approach to characterization of individual carcinogens.[19] An initial qualitative assessment would be made in accordance with the original categorization system. For each particular substance falling into Category I, an assessment would be made of whether current exposure levels constituted a significant risk. Those chemicals considered to pose such a risk would be regulated down to the lowest feasible level, as before.

Of greater import for the Generic Carcinogen Policy was the presidential inauguration several days later of Ronald Reagan. OSHA's cancer policy was listed by the new Administration's Task Force on Regulatory Relief as one of the first 27 regulations to be reconsidered due to their economic costs. On March 27, 1981, the Generic Carcinogen Policy amendments proposed in January were withdrawn in order to permit OSHA to "address alternatives that had not been fully considered."[20] The policy remained in abeyance from then on. In January of 1982, OSHA officially proposed a partial stay on the publishing of the Candidate and Priority Lists of chemicals[21]; this stay was promulgated a year later.[22]

These actions were largely of symbolic importance, since the agency was not pursuing the hazard identification and regulation provisions of the Generic Carcinogen Policy in any case. In August of 1986, the agency announced its intention to propose rulemaking to revise the standard, but did not follow up with any substantive initiative.[23] In February 1987, the Fifth Circuit Court of Appeals dismissed 16 petitions filed by industry and labor groups concerning various provisions of the Generic Carcinogen Policy on the grounds that nothing had happened in six years and hence that it could be considered dormant.[24] Nevertheless the Generic Carcinogen Policy has remained as OSHA's official policy.[25] The National Institute for Occupational Safety and Health (NIOSH) continues to propose lowest feasible level regulations for substances with laboratory evidence of carcinogenicity, under the authority of the Generic Carcinogen Policy.[26,27] This implies that only a relatively simple administrative decision might be sufficient to revive the policy, a matter of concern for industry groups and the Reagan Administration as the election of 1988 approached. The industry-sponsored American Industrial Health Council and the federal Office of Management and Budget (OMB) both recommended that rulemaking be pursued by OSHA to officially excise the offensive parts of the policy.[28,29] The OMB was particularly harsh on OSHA for allowing the timetable of deregulation to slip, declaring its concern for the Reagan "legacy" and noting that "there is only so much time left." Nevertheless, no action was taken on the policy.

The Generic Carcinogen Policy represents OSHA's most aggressive initiative directed at occupational carcinogens. It embodies all the most conservative, health-protective assumptions about the biological process of carcinogenesis. Of greatest importance, it declares that regulation should proceed for those chemicals for which there exists laboratory but not yet epidemiological evidence of cancer effects. The policy weighs the costs and benefits of regulation in a manner strongly oriented toward regulation. OSHA's definition of "feasible" is a highly stringent one. While not unique in its adoption of this concept of feasibility, the Generic Carcinogen Policy stands out in its effort to expeditiously extend the principle to a sweeping array of chemicals.

Substance-Specific Regulations

OSHA continued to promulgate substance-specific standards in the 1980s, but now with even less success than during the previous decade. The battles of the 1970s had produced three legal and institutional barriers to aggressive regulation of occupational carcinogens which continue to haunt the agency to the present.

- First, the Supreme Court's "significant risk" doctrine underlying the dismissal of the original benzene standard requires the agency to perform a detailed quantitative risk assessment for each candidate carcinogen.
- Second, industry concern over the economic costs of compliance contributed to the development of an increasingly rigorous set of requirements for the agency to estimate the economic impact of each proposed regulation. This requirement for economic studies originated with President Gerald Ford,[30] was extended by

President Jimmy Carter,[31] and culminated with President Reagan's requirement for formal cost-benefit analysis.[32] While OSHA's health standards are officially exempt from these cost-benefit requirements due to a Supreme Court ruling on the cotton dust standard,[33] the agency is required to produce rigorous analyses of the expected cost of compliance with its standards.

- Third, the regulatory strategy of the 1970s raised OSHA to a position of high visibility and vulnerability with respect to the Office of Management and Budget, which has interpreted its role as one of reining in a regulatory process gone amok.[34]

Given these impediments and the general anti-regulatory mood of the Reagan Administration, it is perhaps remarkable that OSHA completed any cancer regulations at all. Its record of achievement during the 1980s was quite modest, consisting of leftover business from the previous decade in the form of the revised benzene[35] and asbestos[36] standards plus the promulgation of two new regulations governing ethylene oxide[37] and formaldehyde.[38]

Two features characterize OSHA's new substance-specific regulations. First, they all cover substances for which there exists epidemiological evidence of carcinogenicity in worker populations. While this might charitably be interpreted as beginning at the top of a priority list similar to those produced by the Generic Carcinogen Policy, it actually constitutes a rejection of the guiding principle of the Generic Carcinogen Policy, namely that regulation should proceed based on animal evidence without waiting for epidemiological studies. The Generic Carcinogen Policy was consciously designed to make the transition from epidemiology to animal laboratory evidence as the basis for OSHA regulations; the similarity between the substance-specific regulations of the 1980s and their counterparts in the 1970s eloquently testifies to the failure of this transition. The substance-specific regulations also symbolize the abandonment of OSHA's efforts to provide leadership for the federal regulatory agencies. This role has now reverted to EPA, which has a carcinogen policy that relies heavily on animal evidence.[39] The second salient feature of the substance-specific regulations is their relative stringency. The permissible exposure limit was reduced by 90 percent in the cases of asbestos and benzene, by 98 percent in the case of ethylene oxide, and by 66 percent in the case of formaldehyde.

The Hazard Communication Standard

The most important standard promulgated by OSHA during the 1980s was the Hazard Communication Standard,[40] which requires labeling of hazardous substances and the development of training programs for workers. While less stringent than the Hazard Identification Standard which was proposed by the Carter Administration[41] and withdrawn in the first months of the Reagan Administration,[42] the Hazard Communication Standard is nevertheless sweeping in scope and potentially significant in impact. Although not focused on carcinogens, the Hazard Communication Standard does impose important new obligations on produces and users of these substances.

The Hazard Communication Standard is officially a "performance standard" which delegates to industry management the right and responsibility to decide which substances are to be included in the labeling and training programs. Nevertheless, the Standard does provide a floor of substances that must be included. This includes chemicals listed as carcinogens by the International Agency for Research on Cancer (IARC) of the World Health Organization[43] plus chemicals included in the National Toxicology Program's (NTP) *Annual Report on Carcinogens*.[44] In addition, chemicals must be designated as potential carcinogens and included in worker training programs if they have produced evidence of carcinogenicity in any well-conducted laboratory studies. According to one estimate, these provisions mandate inclusion of 416 substances based on the IARC and NTP documents plus an additional 2,260 substances with laboratory evidence of carcinogenicity but which are not classified by IARC or NTP.[45] This broad scope of substances immediately evokes the Generic Carcinogen Policy, since it places laboratory data front and center in establishing which substances will be classified as potential occupational carcinogens. Less than 1 percent of these substances have strong epidemiological evidence of carcinogenicity in humans.

While similar to the Generic Carcinogen Policy in terms of scope, the Hazard Communication Standard differs markedly in terms of regulatory requirements. It imposes no permissible exposure limits, much less ones that are the lowest feasible short of ruining the industry. Under the terms of the Standard, chemical manufacturers must prepare summaries of available toxicological and epidemiological information, in the form of a Material Safety Data Sheet (MSDS), and provide these to all purchasers of their products. Manufacturers are also responsible for the labeling of containers of hazardous substances; these labels contain brief warnings and refer to the relevant MSDS. All users of covered substances, not just chemical manufacturers, must develop written hazard communication programs that detail how they will comply with the Standard. These users must also organize worker training sessions which cover the information on the labels and MSDS plus proper handling procedures.

The Hazard Communication Standard provides no direct protections but relies upon worker responses to the new information for its ultimate efficacy. In principle, these responses could occur via turnover[46–48] or collective bargaining.[49–51] While modest in the startup costs it imposes on employers, the Hazard Communication standard has the potential to alter management treatment of suspect carcinogens over the long term. In the short term, however, the Standard's effectiveness will be limited by the poor quality of the MSDS and employer training programs currently in place.[52]

The Air Contaminants Standard

A dramatic departure from substance-specific rulemaking, OSHA's 1989 Air Contaminants Standard[53] imposed permissible exposure limits (PELs) on 164 substances not previously regulated and lowered the PELs for an additional 212 substances that had been regulated under the start-up provisions of the original 1970 Occupational Safety and Health Act. Hailed as the centerpiece of a new approach to standard setting, the Air Contaminants Standard has also been severely critized for the procedure used in selecting chemicals and for its choice of exposure limits.

The Air Contaminants Standard is explicitly based upon the Threshold Limit Values (TLVs) of the American Conference of Governmental Industrial Hygienists (ACGIH), a private non-governmental organization, whereas the 1970 OSH Act designates NIOSH as the primary source for recommended standards. The TLVs and the process by which they are

TABLE 1—Number of Substances Classified as Confirmed or Potential Human Carcinogens by Five Scientific Organizations, and Number of Classified as Carcinogens by OSHA's Four Carcinogen Policies

Total Classified	NIOSH 126	ACGIH 49	NTP 162	EPA 181	IARC 376
Generic Carcinogen Policy	126	49	162	181	376
Substance-Specific Regulations	22	16	19	18	20
Hazard Communication Standard	126	49	162	181	376
Air Contaminants Standard	11	9	6	8	9

TABLE 2—Alternative Risk Management Strategies for Occupational Carcinogens

	Permissible Exposure Limit	Labeling and Worker Training	Exposure and Medical Monitoring
Generic Carcinogen Policy	stringent	yes	yes
Substance-Specific Regulations	stringent	yes	yes
Hazard Communication Standard	none	yes	no
Air Contaminants Standard	weak	no	no

established have come under increased scrutiny, partly as a result of OSHA's reliance upon them. Analysis of the minutes of the TLV committee and other documents has uncovered a major but unacknowledged role played by the producers of the chemicals.[54] An analysis of the epidemiological studies referenced as part of the documentation for the TLVs found that the TLVs were generally established at levels commonly reported as prevalent in industry rather than at levels below which no significant adverse health effects were reported.[55] The TLV-based PELs for 98 substances in the Air Contaminants Standard were considered insufficiently protective by NIOSH.[56] Both industry and labor representatives asserted that the PELs mandated by the Air Contaminants Standard may create some problems of compliance for small firms but none for major producers.[57]

The treatment of occupational carcinogens under the Air Contaminants Standard is inconsistent and incomplete. While 78 of the substances covered by the Standard are considered confirmed or potential human carcinogens by NIOSH, ACGIH, and/or NTP, only 11 are regulated under the standard based on cancer risks. The others are regulated based on noncarcinogenic effects. Excluded altogether from the Air Contaminants Standard are 68 substances considered to be confirmed or potential human carcinogens by NIOSH, ACGIH, and/or NTP.[58]

The Air Contaminants Standard makes no consistent distinction between epidemiological and laboratory evidence in regulating substances as carcinogens. Of the 11 substances regulated as carcinogens, six had epidemiological evidence of cancer effects in workers while five had only laboratory evidence of cancer in animals. Of the 67 substances regulated based on noncarcinogenic effects alone, five were considered confirmed human carcinogens by NIOSH, ACGIH, and/or NTP based on epidemiological evidence. Of the 68 substances considered as confirmed or potential carcinogens by the three scientific bodies but excluded altogether from the standard, 11 had epidemiological evidence of cancer in humans.

The Air Contaminants Standard is also inconsistent in the stringency of the exposure limits placed on particular substances. Of the 11 substances designated as carcinogens, seven have PELs set equal to the TLV, three have PELs set lower than the TLV, and one has a PEL set higher than the TLV. OSHA computed the residual risk of cancer (maximum likelihood estimate) for workers exposed to seven of these 11 substances at the new PELs. These residual cancer risks ranged from 0.3 cases per 1,000 workers to 40 per 1,000, with a median of 3.7 and a mean of 10.1. By way of comparison, the maximum likelihood estimates of the residual risk under the substance-specific standards promulgated during the 1980s ranged from 0.0006 per 1,000 for formaldehyde[59] to 1.7 per 1,000 for ethylene oxide,[60] 6.7 per 1,000 for asbestos,[61] and up to 9.5 per 1,000 for benzene.[62]

Conclusion

OSHA's official cancer policy, the 1980 Generic Carcinogen Policy, is comprehensive in the universe of substances covered and stringent in the exposure limitations proposed, but languishes in ignominious neglect. The substance-specific regulations present a politically risk-averse contrast. Eschewing chemicals lacking epidemiological evidence of cancer in humans, they devote enormous resources to documenting significant risk and economic feasibility and impose permissible exposure limits on a very few substances. The 1983 Hazard Communication Standard rivals the Generic Carcinogen Policy in the scope of substances covered but establishes a quite different and less direct form of incentives for management to reduce workplace exposures. The 1989 Air Contaminants Standard rivals the Generic Carcinogen Policy in imposing permissible exposure limits for hundreds of toxic chemicals, but abandons internal consistency with respect to the stringency of the particular limits imposed and external consistency with the recommendations from scientific bodies.

The inconsistencies between OSHA's four carcinogen policies are highlighted on Tables 1 and 2. Table 1 presents the number of substances commonly found in the US workplace environment that are designated as confirmed or potential human carcinogens by five scientific bodies, plus the number of substances on each of these five lists that is covered by each of OSHA's four carcinogen policies.[53] The breadth of the Generic Carcinogen Policy and the Hazard Communication Standard is illustrated by their coverage of all substances listed by the five scientific bodies. In start contrast, OSHA's substance-specific regulations and the Air Contaminants Standard cover only a very small fraction of the listed carcinogens. Most remarkable, perhaps, is the contrast between cancer designations in the Air Contaminants Standard and the designations by NIOSH, the organization to which OSHA officially responds. NIOSH considers 76 of the substances covered by the Air Contaminants Standard to be potential human carcinogens, but OSHA only set 11 permissible exposure limits based on avoidance of cancer risk. The Standard excludes altogether 26 substances considered occupational carcinogens by NIOSH.

Table 2 describes that risk management strategy embodied in each of OSHA's four carcinogen policies. Two of the policies impose stringent permissible exposure limits, one imposes weak limits, and one imposes none. The encouragement of worker self-help initiatives through container labeling and employee training is comprehensive in three policies but absent from the Air Contaminants Standard. Only the Generic Carcinogen Policy and the substance-specific regulations contain provisions for exposure monitoring and medical surveillance.

Progress in society's efforts to reduce the incidence of work-related cancer will only come as part of a general reform of occupational health policy at OSHA. On the political front, the shortest odds are for continued stalemate and perpetuation of the status quo. Among outside observers of OSHA's policies, however, a consensus may be growing as to a way to save generic rulemaking.[13,64] This would involve a major compromise to give up the illusion of imposing exposure limits at the lowest feasible level, in exchange for the ability to rely upon animal evidence of carcinogenicity and hence accelerate the pace of standard-setting. Briefly, OSHA could focus on a wide range of substances for which the available evidence might not survive a "significant risk" judical review but which could pass a less stringent "material impairment of health" review. For these substances, OSHA could impose moderate "best available technology" standards modeled upon EPA experiences with the Clean Water Act.[65] More stringent (lowest feasible) standards would be reserved for the much smaller class of substances that could withstand a "significant risk" review.

This proposal is certain to elicit criticism both from those dissatisfied with the outcome of technology-based standards for environmental regulation and from those attached to the symbolic value of the "lowest feasible" regulatory approach. Technology-based standards provide no guarantee that the residual risks are acceptably low and may create disincentives for industry to develop new control methods that might then be mandated as the "best available."[66] Nevertheless, the technology-based approach retains its pragmatic appeal and has been incorporated for the control of hazardous air pollutants in the revised Clean Air Act.[67] As for critics insistent on OSHA's lowest feasible level strategy, the fate of OSHA's Generic Carcinogen Policy should serve as a note of caution. The real world of power and politics seems to offer stringent limits and narrow scope, as in the substance-specific standards, or broad scope and weak limits, as in the Air Contaminants Standard. What is needed is a mutually consistent and supportive set of standards that deal with each different aspect of the overall risk management strategy. When combined with the Hazard Communication Standard and the proposed generic exposure monitoring and medical surveillance standards,[68,69] an appropriately revised exposure limit policy could provide the foundations for an effective generic approach to controlling occupational carcinogens. ☐

James C. Robinson, PhD
Dalton G. Paxman, PhD

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James C. Robinson and Dalton G. Paxman are with the Center for Occupational and Environmental Health, University of California, Berkeley, CA.

Requests for reprints should be sent to James C. Robinson, PhD, Center for Occupational and Environmental Health, School of Public Health, University of California, Berkeley, CA 94720.

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Health Care Received by American Children: NCHS Report

The vast majority (83 percent) of American children are covered by health insurance, have a regular source of health care (88 percent), and visited a doctor in the past year for routine health care (64 percent) according to a recent report issued by the National Center for Health Statistics. The report cites new evidence, however, that America's minority and poor children receive less care, including preventive services.

The report, *Health Insurance and Medical Care: Health of Our Nation's Children, United States, 1988*, is derived from data from the 1988 National Health Interview Survey on Child Health, conducted by NCHS. Questions on health insurance and sources of medical care were asked for for all 17,110 children in a nationally representative sample of children 17 years of age and under.

The survey shows that Black newborns and infants were 2 to 3 times more likely to have lacked a routine visit for a check-up or any immunization during the first year of life than were White newborns and infants. Hispanic children were twice as likely to have no health insurance coverage, either public or private, as non-Hispanics.

Most children with a regular source of health care had private care; however, 16 percent of those children (8.7 million) received health care from hospital clinics, emergency rooms, walk-in or emergency centers or other clinics. Clinics were the source of care for 47 percent of Black infants, compared to 16 percent of White infants. For Hispanic infants, almost half received health care in clinics.

The report, *Health Insurance and Medical Care: Health of Our Nation's Children, United States, 1988* is available from the National Center for Health Statistics, Centers for Disease Control, US Public Health Service, 6525 Belcrest Road, Hyattsville, MD 20782; (301) 436-8500.

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Commentary

How Blind Is Blind Review?

Alfred Yankauer, MD, MPH

ABSTRACT

Background: No representative surveys of scientific opinion about blind review have been published, and there is very little information on the success of the blinding process. The *American Journal of Public Health* has practiced blind review since 1977.

Methods: In 1989 to 1990 312 of its reviewers were asked to identify author and institution in the manuscript they reviewed, to provide clues to such identification, to express their opinion concerning blind review, and to offer reasons for their opinion.

Results: Reviewers claimed to be able to identify author and/or institution in 47% of the 614 chances offered; identification was incorrect 16% of the time, overall identification correct 39% of the time. Self-referencing was the clue to identification in 62%, personal knowledge in 38% of the cases. If only personal knowledge cases are considered, blinding was successful 83% of the time. Blinding was favored by 75% of the reviewers with most asserting it eliminated bias. Reasons given for opposing blind review included the following: blinding not possible, identification will not influence judgment, and its obverse, identification assists judgment.

Conclusions: For the *American Journal of Public Health* blinding is usually, but not always, successful; and the majority of its reviewers favor current policy. Until more definitive data are in, reviewer preference, which differs from journal to journal, seems the most legitimate guide to journal policy on blind review. (*Am J Public Health*. 1991;81:843–845)

> Some judge of authors' names, not
> words, and then,
> Nor praise nor blame the writings, but
> the men.—*Alexander Pope*

Introduction

Whether or not to blind reviewers to the identity of the authors whose papers they are asked to review has been hotly debated for at least 20 years. In spite of firmly held opinions pro and con, no representative surveys of scientific opinion have been reported, and there is very little information on the prevalence of blinding as a practice or on the success of the blinding process.

Four surveys of the frequency of blinding have been published,[1–4] but none is based on a random sampling of journals from different scientific disciplines, and their size and response rates often leave something to be desired. Their results suggest that the majority of scientific journals do not practice blind review and that blinding may be more common in the social sciences than in the physical and medical sciences.

Small, narrowly focused studies of reviewer opinion about the desirability of blind review report varying proportions in favor of review: 39% with 11% having no opinion (personal communication, Lew Gidez, FASEB), 41% *(Diabetes Care)*,[5] 49% (37% no opinion, *J Neuropathol-Exp Neurol*),[6] and 89% *(Schol Publ)*.[7]

Personal opinions of the success of blinding not based on actual study also vary: 33%,[8] 50% to 60%,[9] and 90%.[10] The reported success of blinding in *J Neuropathol Exp Neurol* was 66% when 55% of all eligible manuscripts were eliminated because of self-referencing,[6] and 73% *(Arch Intern Med)*[11] when self-referencing was partially eliminated.

Reviewer opinions about signing their reviews will be influenced by journal policy. Some journals have encouraged reviewers to sign their reviews arguing that the review process should be completely open. Journals which encourage the practice report a signing rate of only 43% *(Arch Intern Med)*[11] and 58% *(J Lab Clin Med)*.[12] *Diabetes Care* reported that 80% of its reviewers opposed mandatory signing, but 53% favored optional signing.[5] Two other small reviewer surveys report 77% *(Schol Publ)*[7] and 46%[6] (47% no opinion) favoring anonymity for reviewers.

The editorial board of the *American Journal of Public Health (AJPH)* voted to adopt a policy of blind review in 1976. I was skeptical about its value but agreed to survey reviewer opinion. After the first hundred reviewer returns, it was evident that the vast majority of *AJPH* reviewers favored blind review, and the policy was implemented in 1977. In the 13 following years only two of more than 1000 reviewers have refused to review a manuscript because of the blinding policy.

Contributors to *AJPH* are instructed to submit a second face sheet which includes only the title of the paper. These instructions are usually followed. We remove acknowledgments, but make no further effort to remove identifying page headers (when present contrary to instructions) or to change the text or references. We have been aware that a substantial proportion of our manuscripts are not truly blinded because of text allusions and

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Alfred Yankauer is with the Departments of Family and Community Medicine and Pediatrics, University of Massachusetts Medical Center, Worcester, MA 01655. He is the previous editor of this journal.

Requests for reprints should be sent to Dr. Yankauer at the above address.

TABLE 1—Proportion of Correct Identifications of Author and Institution According to "Yes" and "Maybe" Responses

	Author		Institution		Total	
Response Type	%	n	%	n	%	n
Yes	96	(67)	89	(87)	92	(154)
Maybe	70	(66)	80	(66)	75	(132)
Total	83	(133)	85	(153)	84	(286)[b]

[a]Difference in proportion: Yes-Maybe total = 17 (95% correct identifications, 25.5, 6.5).
[b]Both chances taken: 112; Author only taken: 24; Institution only taken: 38.

TABLE 2—Proportion of Correct Identifications of Author and Institutions According to Type of Chance Taken

		Correct Identification (%)	
Type of Chance Taken	N	Author	Institution
Author only	24	83	—
Institution only	38	—	82
Both author and institution	112	83	87
Total	286[a]	83	85

[a]Both chances taken doubled in total.

self-referencing; we so inform reviewers in a covering letter that accompanies the manuscript.

Because of the current interest in blinding I undertook a more extended study in late 1989 and early 1990. Its purpose was to find out how often blinding was successful, to identify the give-away clues, and to determine the opinions of our reviewers about the policy. Our communications to the reviewer were unchanged except for the addition of a questionnaire.

Methods

A short questionnaire was sent to reviewers with consecutive blinded manuscripts. It asked whether author and institution were identifiable (yes/no/maybe) and if so, to record their names and describe the clues that led to their identification. Identification was considered correct if any one of the several authors or institutions represented were listed in the response. A second set of questions asked whether they favored blind review (yes/no) and their reasons for such an opinion. Reviewers were asked to return the form to the editor with their review. They were requested to complete the form only once although they may have reviewed more than one paper during the period of the study.

One of the arguments against blinding is that the review process should be completely open, i.e., if the identity of authors is disclosed, reviewers, in turn, should be willing to sign their reviews. For this reason, persons who were not in favor of blind review were sent another questionnaire. They were asked whether they would like to receive the full face sheet containing the name and institution of the author on the manuscripts sent to them for review (yes/no/no preference) and, if yes, whether they intended then to sign their reviews (yes/no).

Results

The analyses are based on 305 fully completed and 7 partially completed returns (accounting for slight discrepancies in totals). There were no written refusals although 4 blank questionnaires were returned with a review. There were 78 pairs and three triplets (different reviewers for the same manuscript) among the returns. Thus the data reflect responses of 312 reviewers to 228 manuscripts.

There were responses to 614 potential chances to identify author or institution, 286 (47%) of which were taken. Overall 84% of the responses taken correctly identified author or institution (Table 1). "Maybe" responses were less likely than "yes" responses to be correct, but only 25% of the "maybe" responses were incorrect. Author and institution were identified correctly to the same extent even when responses were disaggregated (Table 2). Thus blinding could be considered successful 53% or 61% of the time, depending on whether successful blinding ignores identification or includes only correct identification.

Among those reviewing the same manuscript there was agreement in 108 (90%) of the 120 chances taken in common; in 11 of the 120 chances the agreement consisted of an incorrect identification by both respondents.

Self-referencing (61.8%) and personal knowledge (38.2%) were the two clues given for identification of author and/or institution. In both cases 16% of the identifications were incorrect. Clues given and incorrect identifications were unrelated to opinions concerning blind review.

Overall, blinding was favored by 75% of the respondents with 8% stating they had no preference. The 17% initially opposed to blind review is reduced to 11%, and the proportion with no preference increased to 14% when the follow-up of those opposed to blind review is considered.

The principal reason given for their opinion by those favoring blind review was the elimination of bias (Table 3). The principal reason given for their opinion by those opposing blind review was failure of the attempt to blind (in 2 out of the 14 chances taken by this group, identification was incorrect). Two other reasons concerned with the effect of knowing author identity on the review itself appear to contradict each other.

Of 30 reviewers initially opposed to blind review who, on follow-up, expressed an interest in receiving the manuscript face sheet, only 8 stated they would sign all their reviews; 19 indicated they would not do so, and 3 would do so only in some cases.

Discussion

Blinding of this group of reviewers was somewhat less successful than has been reported by others,[6,11] perhaps because self-referencing was more common. Moreover, in the type of population studies published by *AJPH,* the text must identify the source of the data. However, self-referencing that yields an unfamiliar name may mean nothing to a reviewer. Many reviewers noted that although self-referencing gave away the name, they were unfamiliar with the authors or their work. If all 82 responses that mentioned only self-reference are eliminated, blinding of author could be considered success-

TABLE 3—Reviewers* Stated Reasons for Opinions About Blind Review

Favor Blind Review	Number
Objectivity: judge content, not author	143
Comfort: less pressure, less hesitation to criticize	14
Helps beginning researcher if name is unknown	12
Total	169
Oppose Blind Review	
Blinding does not work	18
Names do not influence my judgment	15
Knowing author's background and track record contributes to judging paper.	14
Total	47

*Ninety-six respondents (31%) gave no reason for opinion.

ful, whether or not identification was correct, 83% of the time. Many reviewers who favored blind review commented that they could often identify the author.

It is possible that the long-standing blind review policy of the *AJPH* may have influenced the opinion of our reviewers. However, as reported elsewhere,[13] the vast majority of these reviewers also review for many other journals, most of which do not have such a blind review policy.

All reports of the success of manuscript blinding reflect a Hawthorne effect whose impact is unknown: reviewers may seek out author and institutional identity more vigorously than in their usual review because the question is asked. Evidence for this constraint on the interpretation of these data can be found in this study: in 11 of the 120 chances taken in common, 1 reviewer identified author and/or institution correctly, giving as the reason self-reference, whereas the other reviewer of the same manuscript claimed to be unable to make any identification.

Most of those opposed to blind review would not sign their reviews, throwing doubt on the feasibility of a truly "open" review process. As an editor I, too, would be opposed to having the reviewers sign their reviews because their signing would intrude a third person into the decision process. Those who favored blind review were not asked whether they would sign their reviews.

This survey was not designed to assess the effect of blinding on the content of the review or the reviewers' recommendations to the editor. Nevertheless, the findings can be interpreted as evidence both for and against blinding. Against blinding is the fact that in about half the manuscripts some identification was claimed. Thus if identification biases the reviewer, it does so for some authors but not for others, affecting the equity of the review process; an appreciable number of cases may be falsely identified. In favor of blinding is the fact that a small number of reviewers opposed to blind review apparently felt that the track record and background of authors were factors that could influence their judgment about a paper. This may be reasonable for the rare opinion piece, but it implies a possible bias in the case of a research paper that should be judged solely on its merits.

The vast majority of *AJPH* reviewers favoring blind review contrasts with a recent report of a reader survey of the Federation of American Societies for Experimental Biology (FASEB) journal; most of the respondents were also reviewers (Lew Gidez, personal communication). Response rate for this survey was 45%, but there were 871 respondents. Half of the respondents were opposed to blind review compared with 17% or less of the *AJPH* reviewers. One can only speculate about reasons for this difference of opinion; the fact that the *FASEB J* reviewers are closer to the physical sciences and that *AJPH* reviewers are closer to the social sciences may account for some of this difference.

Although blind review may produce a somewhat kinder and more complete review,[11] I doubt that it can remove bias where negative or positive reviewer bias is associated with reviewer knowledge of author identity. I believe one must rely on the reviewers themselves to be frank about their own biases. Many reviewers return papers with a note recusing themselves for a variety of reasons. Moreover, it is the editor, not the reviewer, who makes the decision on a manuscript, and the editor cannot be blinded.

The fact that the majority of *AJPH* reviewers prefer to continue the blind review policy seems sufficient reason to continue it. Until more definitive data are in, reviewer preference, which differs from journal to journal, seems the most legitimate guide to the policy of journals on blind review. ☐

Acknowledgments

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Tobacco Smoking as a Risk Factor for Colon Polyps

Shelia Hoar Zahm, ScD, Pierluigi Cocco, MD, and Aaron Blair, PhD

ABSTRACT

Background: Data from a cancer screening project among pattern workers were used to evaluate the association between tobacco smoking and prevalence of colon polyps.

Methods: From 1981–1983, 549 White men were examined by flexible sigmoidoscopy and completed self-administered questionnaires including smoking histories.

Results: One or more colon polyps were detected in 76 men. Standardized prevalence rates (SPR) for polyps increased by smoking category (never smoked = 0.094; ex-smokers = 0.118, current smokers = 0.214) and by cigarettes per day, years of smoking, and pack-years among both current and ex-smokers. Both adenomatous and hyperplastic polyps showed an association with smoking while other types of polyps and polyps with unspecified histology did not. The risk associated with smoking was greater for polyps greater than one centimeter in diameter. An interaction with occupational exposures was suggested by a greater increase in the SPR for polyps among current smokers employed as pattern makers for more than 10 years than among current smokers similarly employed for 10 years or less.

Conclusions: Since at least some colon polyps are considered precursor lesions to colon cancer, one of the most common cancers in the United States, this report suggests that the possible link between colon polyps and smoking deserves further evaluation. (*Am J Public Health*. 1991; 81:846–849)

Introduction

Most epidemiologic studies of colon cancer have suggested that there is no strong or moderate elevation in risk of cancer of the colon among smokers. [1,2] Whether there is a slight increase in risk, perhaps limited to a subgroup of susceptible people, is not known, although some recent work suggests this possibility.[3–6] In addition, an association has been noted between smoking and the presence of adenomas and hyperplastic polyps of the colon in men and women.[7,8] Since adenomatous polyps are known to be precursors of colon cancer,[9] the possibility that smoking influences the development of colon polyps deserves further evaluation. The availability of the data from a colon cancer screening project among pattern makers allowed us to evaluate the association between tobacco smoking and prevalence of colon polyps.

Methods

Pattern makers construct prototypes for mass-produced items, such as automobiles and agricultural implements. They can be exposed to wood, metals, fiberglass, plastics, epoxy resins, welding or soldering fumes, cutting oils, solvents, and spray paints. Three partially overlapping epidemiologic studies reported excess colorectal cancer in pattern makers[10–12] and led to cancer screening programs of pattern makers employed by General Motors Corporation and Ford Motor Company[13] and of members of the Pattern Makers League of North America who worked independently or for other corporations.[13] Data from the screening program of the non-General Motors members of the Pattern Makers League employed other than at General Motors Corporation or Ford Motor Company are the subject of this report. The background of this project has been described in more detail elsewhere.[14,15]

From 1981 through 1983, 1,465 White men from 16 factories were examined by flexible sigmoidoscopy and completed a self-administered questionnaire collecting information on sex, age, race, marital status, education, cancer symptoms, respiratory function, medical history, occupational history, and a few dietary constituents including alcohol but not total fat. The participation rate was 48 percent, a low rate due in part to the economic depression that affected the United States automotive industry during the screening program.[12] The unemployment rate among League members rose from 3 percent to 53 percent during the program with many unemployed pattern makers becoming ineligible for the medical insurance coverage that supported the screening.

Men from 11 of the factories (N = 549) also provided smoking histories and form the basis of this report. There is no evidence that the smoking histories, exposures, or polyp rate of the men at the remaining five factories where the workers did not provide smoking histories differed from that at the 11 other factories. The participation rates at the 11 factories where workers provided smoking histo-

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From the Occupational Studies Section, National Cancer Institute. Address reprint requests to Shelia Hoar Zahm, ScD, Epidemiologist, Occupational Studies Section, NCI, Executive Plaza North, Room 418, Rockville, MD 20892. Dr. Blair is Section Chief at NCI. Dr. Cocco was on a research training fellowship from Associazione Italiana per la Ricerca sul Cancro, Milano, Italy; he is currently with Istituto di Medicina del Lavora, Cagliari, Italy.
This paper, submitted to the Journal April 6, 1990, was revised and accepted for publication December 20, 1990.

TABLE 1—Characteristics of the Pattern Maker Population

Characteristics	Number	%
Total	549	
Age (years)		
<= 40	191	35
41–55	221	40
56+	137	25
Duration of employment (years)		
<= 10	440	80
11–20	59	11
21+	50	9
Employment Status		
Active	496	90
Inactive	25	5
Retired	15	3
Other	9	2
Unknown	4	1
Diagnosis		
Cancer	3	1
Polyps	76	14
Other gastrointestinal disease	54	10
No gastrointestinal disease	411	75
Unknown	5	1

ries (47 percent) and at the remaining five factories (50 percent) were similar. The age, duration of employment, employment status (e.g. active, inactive, and retired), and screening diagnoses of the study subjects are presented in Table 1.

The screening program identified three men with colon cancer, 76 men with one or more colon polyps, 54 men with other gastrointestinal disease, and 411 men with no evidence of any gastrointestinal disease. Five men had unknown sigmoidoscope results. Thirty-one of the 76 polyp cases had adenomatous polyps, 24 had hyperplastic polyps, and the remainder had other or not otherwise specified polyps.

Cigarettes per day, years of smoking and pack-years were used as indices of exposure to tobacco smoking for each current or ex-smoker. Three men without colon cancer or polyps had unknown smoking histories and were excluded from the analyses. Several subjects had incomplete smoking histories (e.g. ex-smokers with unknown number of cigarettes per day) and were excluded from some analyses when necessary.

We computed standardized prevalence rates (SPRs) for each category of smoking habit, adjusted for age (< = 40 years, 41–55 years, and 56+ years) and alcohol use (daily consumption of beer, wine, or hard liquor: no, yes, unknown), using the age and alcohol distribution of the entire study population participating in the screening as the standard. Smaller categories of age were not necessary for adequate adjustment for age. The measure of association was the risk ratio (RR). Maximum likelihood estimates, adjusted for the age and alcohol distribution of the smokers and the non-smokers among the men from the 11 factories who provided smoking histories, 95 percent confidence intervals (CI), and Chi tests for trend were obtained using EPITOME, personal computer statistical software under development by the Epidemiologic Methods Section, Biometry Branch, of the National Cancer Institute.[16]

Results

The age- and alcohol-adjusted prevalence rate for colon polyps among men who had ever smoked (0.150) was significantly higher than the rate among nonsmokers (0.094) (RR = 1.7, CI = 1.0, 2.9) (Table 2). The prevalence rose by smoking category (ex-smokers = 0.118; current smokers = 0.214; p for trend = 0.008) and by average number of cigarettes per day, years of cigarette smoking, and pack-years among ex-smokers. The increases in cigarettes per day, years, and pack-years of smoking were also significant, but less consistent, among current smokers.

Comparing current smokers to men who never smoked, risks for adenomatous (RR = 2.7, CI = 1.0, 7.1) and hyperplastic polyps (RR = 3.0, CI = 1.0, 8.7) were higher than risk for other or not specified polyps (RR = 1.0, CI = 0.3, 4.1) (Table 3).

The relative risk for smoking also varied by size and number of polyps. Among current smokers, the RR for polyps less than 1.0 centimeter in diameter was 1.5 (CI = 0.7, 3.1) in comparison with a RR of 2.6 (CI = 1.0, 7.0) for larger polyps. Men who never smoked and ex-smokers were at greater risk for only one polyp than for two or more polyps. Among current smokers, the RR was similar for only one (RR = 1.5) and more than one polyp (RR = 1.6).

Among current smokers, risk for polyps was greater in men employed as pattern makers for more than 10 years (RR = 3.0, CI = 1.3, 7.0) than in men employed 10 or fewer years (RR = 1.4, CI = 0.7, 2.9) (Table 4). Greater risks of polyps among the smokers employed more than 10 years were observed at almost all levels of smoking intensity and duration in comparison to the smokers employed 10 or less years. The higher risks in the longer employed workers were not due to a disproportionate number of heavy or long-term smokers in the group in comparison to the smokers employed 10 or less years. There was little difference by duration of employment among men who never smoked or were ex-smokers.

Discussion

The results presented in this report support the previous findings[7,8] of an association between tobacco smoking and risk of colon polyps. Risk increased significantly and consistently with smoking history category, cigarettes per day, years of smoking, pack-years, and diameter of polyps. Smoking increased the risk for hyperplastic and adenomatous polyps, which are considered the primary precursor lesions of colon carcinoma,[1,9,17] more than the risk for other histologic types. The risk of colon polyps among current smokers was greater in men employed as pattern makers more than 10 years than in men employed for shorter periods of time.

Cigarette smoking is known to be strongly associated with coffee consumption[18] and may be associated with other dietary factors, such as fat intake[18] and alcohol consumption,[19–21] which influence the development of colon polyps and cancer. Analyses were adjusted for age and alcohol consumption, but detailed dietary histories were not available in this study to evaluate other possible confounders. Even the adjustment for alcohol, based on only a crude classification, may have been inadequate to completely account for its possible confounding.

The present study was based on prevalent colon polyps detected in a screening program. Because prevalence reflects incidence and duration, it is not clear if the association with smoking is related to duration, rather than incidence, of colon polyps. The increase of risk with the diameter of polyps is consistent with an association with duration; however, differences in growth rates may be responsible. In addition, risk of multiple polyps was not greater than risk of single polyps among smokers.

No information is available on previous colonoscopies among the study subjects; however, the subjects were asked whether a doctor had ever told them they "had a polyp (growth) in the large intestine (colon, rectum, bowel)." The same percentage of smokers (4 percent) and nonsmokers (4 percent) reported a previous diagnosis of polyp. Therefore, there is

TABLE 2—Standardized Prevalence Rates (SPR) and Risk Ratios (RR) for Colon Polyps by Smoking History*

	N	SPR*	RR**	95% CI++
Never smoked	17/185	0.094	1.0	
Ever smoked	59/361	0.150	1.7	1.0, 2.9
Ex-smokers	31/229	0.118	1.4	0.8, 2.5
Cig/day: ≤19	6/74	0.076	0.9	0.4, 2.2
20–39	16/107	0.142	1.5	0.8, 2.9
40+	6/30	0.192	2.5	1.1, 5.8
Test for trend p = 0.0146				
Years of smoking: ≤10	7/92	0.073	1.2	0.5, 2.7
11–25	8/84	0.118	1.1	0.5, 2.4
26+	14/37	0.169	3.0	1.4, 6.4
Test for trend p = 0.0028				
Pack-years: ≤20	9/126	0.069	0.9	0.4, 2.0
21–40	6/42	0.081	2.1	0.9, 4.7
41+	13/43	0.253	3.3	1.6, 6.9
Test for trend p = 0.0013				
Current smokers	28/132	0.214	2.2	1.2, 4.1
Cig/day: ≤19	7/29	0.276	5.5	2.5, 12.1
20–39	12/76	0.146	1.5	0.7, 3.3
40+	7/20	0.330	5.7	2.6, 12.9
Test for trend p = 0.0035				
Years of smoking: ≤10	0/13	0.000	—	
11–25	6/51	0.074	3.3	1.3, 8.3
26+	21/62	0.192	2.8	1.4, 5.5
Test for trend p = 0.0006				
Pack-years: ≤20	3/48	0.069	2.4	0.9, 6.4
21–40	12/41	0.307	4.5	2.2, 9.4
41+	12/41	0.174	2.1	1.0, 4.5
Test for trend p = 0.0014				

*Three subjects with unknown smoking histories were excluded.

+SPR = Standardized prevalence rate, adjusted using the age and alcohol distribution of the entire study population participating in the screening program as the standard population.

**RR = Risk ratio, maximum likelihood estimate, adjusted for the age and alcohol distribution of the smokers and nonsmokers among the men from the 11 factories that provided smoking histories.

++95% confidence interval. Reference category for all comparisons is men who never smoked.

TABLE 3—Rate Ratios (RR) for Colon Polyps by Smoking History and Histologic Type, Diameter, and Number of Polyps

	Never Smoked (185)			Ex-Smoker (229)			Current Smoker (132)		
	N	RR	95% CI*	N	RR	95% CI	N	RR	95% CI
Histology									
Adenomatous	7	1.0		12	1.2	0.5, 2.7	12	2.7	1.0, 7.1
Hyperplastic	5	1.0		7	0.9	0.3, 2.4	12	3.0	1.0, 8.7
Other and NOS	5	1.0		12	1.6	0.6, 4.1	4	1.0	0.3, 4.1
Diameter of Polyp+									
<1 cm	14	1.0		20	1.1	0.5, 2.1	17	1.5	0.7, 3.1
≥1 cm	3	0.7	0.3, 1.9	11	1.0	0.4, 2.4	11	2.6	1.0, 7.0
*Number of Polyps***									
1 only	12	1.0		21	1.3	0.7, 2.6	14	1.5	0.7, 3.3
2 or more	5	0.4	0.1, 1.2	10	0.6	0.2, 1.3	14	1.6	0.7, 3.5

*95% confidence intervals, adjusted for age and alcohol.

+Reference category is the men who never smoked with polyps of less than 1 cm diameter.

**Reference category is the men who never smoked with only one polyp.

no evidence that the nonsmokers had more previous diagnostic tests or polypectomies than the smokers. The lower prevalence of polyps observed in nonsmokers in the current screening program does not appear related to differences by smoking status in participation in previous screening programs.

It is difficult to determine if pattern makers as a group have an increased prevalence of colon polyps in comparison to the general population. There are few published reports on the results of colon screening among asymptomatic subjects using flexible sigmoidoscopy. Most reports are based on subjects presenting with symptoms for a gastroenterology evaluation and, until recently, used the shorter rigid sigmoidoscope. Age-specific prevalence data and information on the size of the polyps detected are necessary for unbiased comparisons since the prevalence of polyps increases dramatically with age and the minimum diameter of reported polyps affects prevalence estimates. An earlier report on this population of pattern makers[14] concluded that the prevalence of polyps was not excessive in comparison with four other asymptomatic populations screened by flexible sigmoidoscopy. Another evaluation of pattern makers,[22] which included some of the same subjects as the Hoar, *et al*, report,[14] concluded there was an excess of polyps in the pattern makers; however, the polyp size may not have been considered adequately.

The observed association between smoking and colon polyps, in particular adenomatous polyps whose progression to cancer has been reported,[9,17,21] suggest that colon cancer may also be affected by use of this product. Colon cancer is not generally considered a tobacco-related malignancy; however, the literature is inconsistent. Most published studies of colon cancer have not provided information on smoking. Slight deficits of colorectal cancer among smokers have been reported in the American Cancer Society survey,[23] the Third National Cancer Survey,[19] and a recent study from Italy.[24] A significant inverse association has been reported among women by Sandler, *et al*,[3] which, they speculate, may be due to an antiestrogenic effect of smoking. On the other hand, Sandler, *et al*,[3] reported an increased rate of colon cancer in male smokers ages 25–44, (nonsmokers and passive smokers: 5.5/100,000 ; ever smokers: 19.0/100,000). Slight excesses of colorectal cancer occurred among male British doctors,[4] US veterans,[5] and, for rectal cancer, female British doctors.[6]

Not all colon polyps result in colon cancer. Although some polyps are considered precancerous lesions, some of them disappeared in a follow-up study.[25] The risk factors for colon cancer and polyps may differ or act with different strength.[26] Smoking may interact with other factors, as suggested in this study by the higher risk of polyps in long-term employees.

TABLE 4—Rate Ratios (RR) for Colon Polyps by Smoking History and Duration of Employment as a Pattern Maker

	Employed ≤10 years			Employed 11+ years		
Smoking History	N	RR	95% CI[+]	N	RR	95% CI
Never smoked	15/144	1.0		2/39	1.1	0.4, 3.0
Ex-smokers	27/183	1.2	0.6, 2.3	4/43	0.9	0.3, 3.0
Current smokers	20/103	1.4	0.7, 2.9	8/26	3.0	1.3, 7.0

[*] Reference category for all comparisons is the men who never smoked and were employed as pattern makers for ≤10 years.

[+] 95% confidence interval, adjusted for age and alcohol.

An explanation for the findings of an association between colon polyps and smoking and the inconsistent relationship between smoking and colon cancer in the published literature is not entirely apparent to us but, may be due to population variations in the distribution of non-, ex-, and current smokers and in the percent of colon cancers with adenomatous polyp precursors. In the current study, there were 185 nonsmokers, 229 ex-smokers, and 132 current smokers whose risk ratios for adenomatous polyps were 1.0, 1.2, and 2.7, respectively. Previous reports in the literature suggest that approximately 30 percent of colon cancers have adenomatous polyp precursors.[27,28] Assuming that this 30 percent has the same smoking distribution and risks as the adenomatous polyp cases and that the remaining 70 percent of colon cancer cases have no risk associated with smoking (risk ratio = 1.0), a weighted overall risk estimate would be 1.15. Slight changes in any of these parameters would result in the inconsistent overall risk estimates for colon cancer and smoking observed in literature, particularly in the presence of residual confounding by dietary and other risk factors. Given the high incidence of colon cancer in the United States, the association deserves further evaluation. ☐

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Passive Smoking During the First Year of Life

Robert A. Greenberg, MD, MSPH, Karl E. Bauman, PhD, Victor J. Strecher, PhD, MPH, Lynette L. Keyes, MS, MPH, Lucinda H. Glover, MPH, Nancy J. Haley, PhD, Helen C. Stedman, BS, and Frank A. Loda, MD

ABSTRACT

Background and Purpose: A description of passive smoking during the first year of life might assist planning preventive efforts.

Methods: Changes in the ecology of passive smoking were investigated in a sample of infants in central North Carolina followed from birth to one year of age.

Results: The prevalence of tobacco smoke absorption, indicated by excretion of cotinine, increased from 53 percent to 77 percent (95% CI of difference: 14, 35) during the first year of life. Most infants (92 percent) excreting cotinine at three weeks of age were also excreting it at one year. Moreover, 61 percent of infants not excreting cotinine at age three weeks were excreting it at one year. This increase reflected an increased exposure to household and, particularly, nonhousehold sources of smoke; the proportion of infants exposed to nonhousehold smokers increased from 14 percent to 36 percent.

Conclusions: These findings suggest that prevention of the onset of passive smoking should begin very early. (*Am J Public Health.* 1991; 81:850-853)

Introduction

Recognition of the adverse health effects of passive smoking in childhood, particularly during the first year of life,[1-3] has led to efforts to reduce exposure to environmental tobacco smoke.[4] To better inform such efforts, we previously reported the frequency of passive smoking and the circumstances in which it occurred in healthy neonates in central North Carolina.[5] The observations from that research were extended by a one year follow-up of a sample of the original study infants. This paper reports the changes in tobacco smoke exposure and absorption during the first year of life, identifies selected correlates of these changes, and considers the implications of findings for prevention of passive smoking during infancy.

Methods

Study Population

The infants were enrolled at birth, from April 1986 to May 1987, in an experimental study of a program to reduce passive smoking and respiratory illness. To be eligible for enrollment, an infant had to have no significant postnatal problems, a birth weight of at least 2,000 gm, and reside in Alamance County or Chatham County in central North Carolina. Infants in the previous neonatal study[5] were recruited from three hospitals where 433 mothers (40 percent of the 1,096 eligible) gave informed consent. Following data collection for those 433 infants, 216 infants began the passive smoking reduction program to which they had been randomly assigned and 217 remained in the control group. Sixty-five infants in the control group were lost to follow-up by one year of age. The remainder of the control group (n = 152) constituted the study sample for the present paper.

The mean number of adults living in study households was 2.3 at both enrollment and at one year. Seventy-five percent of the infants were living in two-parent families as neonates and 70 percent at one year.

Table 1 describes the study sample. There were no significant differences between the 152 infants in the study sample and the remainder of the 1,096 infants who were eligible for enrollment in the gender of the infant, mother's age, mother's race, and smoking by the mother. The study population, however, included significantly more heads of household with more than a high school education (45 percent) than the remainder of the eligible population (34 percent).

Data Collection Procedures

Data collectors visited the homes of enrolled infants when the infants were approximately three weeks of age [mean age (SD): 18 (seven) days, range: nine to 51 days] and before the family's assignment in the intervention trial. The follow-up visit for the data reported in this study was when the infants were approximately one year old [mean age (SD): 12.3 (0.6)

From the University of North Carolina at Chapel Hill (Greenberg, Bauman, Strecher, Keyes, Glover, Stedman, Loda) and the American Health Foundation, Valhalla, NY (Haley). Address reprint requests to Robert A. Greenberg, MD, MSPH, Division of Community Pediatrics, CB #7225, Medical School Wing C, University of North Carolina, Chapel Hill, NC 27599-7225. This paper, submitted to the Journal April 23, 1990, was revised and accepted for publication September 20, 1990.

TABLE 1—Description of Study Sample (N = 152)

Variables	Percent or Mean
Infants	
Female	49%
Male	52%
Mother	
Race	
White	74%
Black	26%
Age	
Mean (SD)	25.6 (5.2) years
Smoking behavior	
Smokers	26%
Mean (SD)[a]	13.7 (8.7) cigarettes per day
Household Heads	
Education	
Less than High School Graduate	18%
High School Graduate only	38%
More than High School Graduate	45%

[a] Smokers only.

months, range: 11.5 to 14.9 months]. Before beginning the interview, the data collector placed a urine collection bag on the infant and explained to the mother that the urine would be analyzed for tobacco smoke products to determine whether the baby had been exposed to smoke. The mother was then asked to describe the smoking habits of all household members and any recent exposure of the infant to tobacco smoke.

Definitions of Variables

An infant was considered to have *contact with smokers* if the infant lived in the same household as a smoker or was in the same place, at the same time, as a nonhousehold smoker, either in the house or elsewhere. A smoker was anyone who smoked at least one cigarette, cigar, or pipeful per day.

Exposure to tobacco smoke was considered to occur when smoke was produced in the infant's presence, that is, in the same room or vehicle as the infant. The amount of exposure was expressed as cigarettes per week smoked in the infant's presence. (See Appendix for method of measurement.)

Smoke absorption was defined as sufficient airway contact with tobacco smoke to result in absorption of nicotine. It was measured by the urinary concentration of cotinine, a major metabolite of nicotine.[2]

Urine cotinine and our questionnaire measure of tobacco smoke exposure are related, but different measures of passive smoking. We have reported analyses of this relationship in infants elsewhere.[5,6] (See Appendix for further discussion of measures of passive smoking.)

Laboratory Procedures

The urine samples were frozen and shipped to the American Health Foundation, in Valhalla, NY. Researchers there analyzed each specimen without knowing the exposure status of the infant.

Cotinine concentration was quantitated by radioimmunoassay with a modification of the method originally described by Langone, *et al.*[7] This protocol uses specific antisera produced by rabbits and has interassay and intraassay variations of 8 percent, with a sensitivity of 1 ng/ml (6 nmol/L). Levels less than 2 ng/ml (12 nmol/L) were reported as below limits of reliable detection.

Creatinine was measured by dry chemistry methods on a Kodak Ektachem 400 analyzer (Eastman Kodak Company, Rochester, NY). The cotinine was divided by the creatinine concentration to adjust the data to reflect the concentration of the urine sample.

TABLE 2—Sources of Exposure to Tobacco Smoke

	Percent of Infants (N = 152)			
Source of Exposure	3 weeks %	1 year %	% Difference (95% CI)	% Change
Any source	39	63	23[a] (14, 32)	+62[a]
Mother	18	29	11 (5, 16)	+61
Nonmaternal household members	24	35	11 (4, 19)	+46
Nonhousehold smokers	14	36	22 (12, 31)	+157
More than one source	15	33	18 (10, 26)	+120

[a] Because of rounding, number does not exactly equal calculations made from table.

Infants of the nine mothers who smoked and breast-fed were eliminated from all analyses that used cotinine as a continuous variable because of the large amount of nicotine and cotinine absorbed through breast-feeding, relative to the amount absorbed through the airways.[8]

Statistical Methods

Distributions of categorical variables were compared by the chi-square statistic or the Fisher's Exact Test. The two-tailed Student t statistic was used to compare means of continuous variables for two groups.

McNemar's test was used to assess changes between three weeks and one year of age in the proportion of infants exposed to tobacco smoke from different sources and the proportion excreting cotinine.

Analyses of cotinine comparing smoke absorption at the two ages involved only the 114 infants who provided urine specimens at both data collection visits.

Results

Exposure to Tobacco Smoke

Between three weeks and one year of age, the proportion of infants reportedly exposed to tobacco smoke from any source increased from 39 percent to 63 percent (Table 2). The proportion of infants exposed to tobacco smoke produced by their mothers, by nonmaternal household members, by nonhousehold smokers, and by more than one of these sources, increased significantly during the first year of life. The greatest increase (157 percent) was in the proportion of infants exposed to smoke produced by nonhousehold smokers. Very similar results were found when the data analyses for Table 2 were repeated for just the 114 infants who provided urine specimens for cotinine analysis.

One infant was in day care at 3 weeks of age. By one year, 90 (58 percent) of the 152 study infants were in day care and 17 (19 percent) of those in day care were exposed there. Only six infants had their only exposure in day care.

The median amount of exposure for the 39 percent of infants who were exposed to any source of smoke at three weeks was 36 cigarettes per week per infant, with a range from two to 250 cigarettes per week. When these exposed three-week-old infants became one-year-olds the median amount of exposure was still 36 and the range was zero to 343 cigarettes per week.

TABLE 3—Distribution of Infants by Smoke Absorption Status at Age One Year Based on Status at Age 3 Weeks (N = 114[a])

	Smoke Absorption Status	
	Age 3 weeks (N)	Age 1 year (N)
Cotinine in urine	+(60)	+(55) 92% − (5) 8%
No cotinine in urine	−(54)	−(21) 39% +(33) 61%

[a]Infants who provided urine samples at both ages.

Among exposed three-week-old infants, the mother contributed, on the average, 34 percent of the cigarettes per week to which an infant was exposed. All household members other than the mother contributed 42 percent and nonhousehold smokers contributed, on the average, 23 percent to the total amount of an infant's exposure. Among exposed one-year-old infants, the mother contributed, on the average, 33 percent; nonmaternal household members, 38 percent; and nonhousehold smokers, 28 percent.

Absorption of Tobacco Smoke

At three weeks of age, 53 percent of the infants had cotinine in their urine. By one year, the proportion excreting cotinine increased to 77 percent (95% CI of difference: 14, 35).

For the infants excreting cotinine at three weeks, the median cotinine concentration was 79 ng/mg of creatinine (51 nmol/mmol) with a range of 9–643 ng/mg (6–413 nmol/mmol). The median among the same infants at one year was 127 ng/mg of creatinine (82 nmol/mmol), range 0–714 ng/mg (0–458 nmol/mmol). Among individual infants, the difference between the urine concentrations at the two ages varied considerably. Although the mean difference was only 11 ng/mg of creatinine (7 nmol/mmol), the 95% CI of the difference was wide: −40 ng/mg (−26 nmol/mmol), 62 ng/mg (40 nmol/mmol).

Table 3 contains the distribution of infants by smoke absorption status at one year, according to smoke absorption status at three weeks. Almost all infants (92 percent) absorbing tobacco smoke at three weeks were also absorbing smoke at one year. In contrast, only 39 percent of infants who were not absorbing smoke at three weeks were still not absorbing smoke at one year. We compared the characteristics of these 21 infants with the 33 infants who also were not absorbing smoke at three weeks, but who became tobacco smoke absorbers by one year. The latter infants were more likely to live in a household with smokers, 30 percent compared with only 5 percent of infants who continued to not absorb smoke at one year. The two groups of infants did not differ by educational level of the head of household, race, or the proportion attending day care at one year.

Discussion

Although passive smoking was already common at three weeks of age, there was a large increase in the prevalence of passive smoking by one year, an increase found in both the proportion of infants exposed to smoke and in the proportion absorbing nicotine. The large number of infants eligible for enrollment who were not in the study sample may limit the generalizability of our results. For example, because the heads of household of infants not included in the study had less education and were therefore more likely to smoke, the extent of passive smoking among study infants may be an underestimate for infants in the entire community.

Our analysis of smoking by household members suggests that the increase in prevalence of passive smoking might be reduced by urging smokers who are already in contact with infants, but not smoking in their presence, to continue to refrain after the newborn period. Considering the large increase during the first year of life in the proportion of infants exposed to smoke produced by nonhousehold smokers, parents should be encouraged to limit contact between their infants and nonhousehold smokers. Mothers may be unaware of all contacts their infants have with smokers, however: 24 percent of the three-week-old and 49 percent of the one-year-old infants in our study were reported to have no contact with smokers but had cotinine in their urine.

Our findings support vigorous early and continued efforts to prevent infant passive smoking. The larger study from which the data reported here are derived is evaluating one such effort.[9] □

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APPENDIX

Measures of Passive Smoking

Data to measure exposure to tobacco smoke were gathered for both "a typical day last week" and "during the last weekend." The mother identified all the places where the infant spent time, including rooms in the house, at other homes, in cars, and in other places outside the home. For each place, the mother reported all smokers who entered while the infant was there. She also reported whether they smoked, as well as what, and how much they smoked. The amount of exposure was calculated by multiplying the number of cigarettes smoked in the infant's presence during "a typical day" of the previous week by five, and adding the number of cigarettes smoked in the infant's presence during the previous weekend. A cigar or pipeful was considered equivalent to one cigarette. Infants who were cared for more than four hours a week away from their home were defined as in day care. If the mother reported that the day-care workers smoked inside where the children were, then the infants were considered exposed to tobacco smoke. The amount of exposure in day care was not determined.

Both the interview measure of exposure to tobacco smoke and urine cotinine were used as measures of passive smoking because of their different strengths and limitations.[6,10] The interview measure estimates the amount, source, and location of exposure during the previous week, taking into account potentially different smoking habits during the week and the weekend. (We requested a report of exposure during a "typical" weekday because we were interested in a reflection of chronic exposure that was not influenced by unusual experiences during the previous week. However, a possible limitation of using only a "typical" weekday is that exposure in some infants may vary considerably from weekday to weekday.) The questionnaire results are limited by being subject to the memory of the mother and to biased reporting.

The presence of cotinine in urine is an accurate indication that nicotine has been absorbed during the 48 to 72 hours prior to specimen collection, but provides no information about the exposure leading to absorption or about past absorption. The concentration of urine cotinine varies considerably among those exposed to the same amount of nicotine,[11] and the possibility of differences in the same infant at different ages in the pharmacokinetics and metabolism of nicotine remains to be explored. Also, nicotine is only one of thousands of compounds in tobacco smoke and is probably not responsible for the health effects of passive smoking.[1]

Among the study infants who provided urine specimens at three weeks and one year of age, we found the following relationships between exposure to tobacco smoke (cigarettes per week smoked near the infant) and absorption of nicotine (urine cotinine concentration). The Spearman rank correlation at three weeks was 0.50 (95% CI: 0.35, 0.63) and at one year, 0.68 (95% CI: 0.56, 0.77). Among the 60 infants excreting cotinine at three weeks, 60 percent were reportedly exposed during the previous week; whereas among the 54 infants with no cotinine in their urine at three weeks, 20 percent were exposed. Among the 88 infants excreting cotinine at one year of age, 69 percent were exposed during the previous week; whereas among the 26 infants not excreting cotinine at one year, 27 percent were exposed.

Are the Correlates of Cigarette Smoking Initiation Different for Black and White Adolescents?

ABSTRACT

Background: Studies of adolescent smoking suggest that the causes of smoking initiation may differ for Blacks and Whites.

Methods: Correlates of smoking initiation were examined among 1,277 nonsmokers, ages 12–14, who completed questionnaires in their homes. The analyses examined relationships between smoking initiation and 11 explanatory variables using logistic regression with the combined sample and with Black and White samples.

Results: Over two years, 24 percent of Whites and 14 percent of Blacks started to smoke. Whites were more likely to start smoking at age 12 and Blacks at age 14. Having a best friend who smoked increased the odds of initiating smoking over twofold for Whites but had no effect on the odds of smoking for Blacks.

Conclusions: Whites initiate smoking earlier than Blacks and are more likely to be influenced by friend behavior. (*Am J Public Health*. 1991; 81:854–858)

Sandra W. Headen, PhD, Karl E. Bauman, PhD, Glenn D. Deane, MA, and Gary G. Koch, PhD

Introduction

Most research on adolescent smoking has involved only White subjects or has not reported findings separately for Blacks and Whites.[1–9] The findings of these studies typically have been applied to both races, but this practice may lead to erroneous conclusions about the behavior of Blacks. The several studies that compared Blacks and Whites found both similar and different correlates of smoking. For example, smoking initiation by both Blacks and Whites has been linked to age, smoking by siblings and friends, perceived availability of cigarettes, intentions to smoke, and ease of refusing cigarettes; risk-taking has been associated with smoking for Blacks but not Whites; parental smoking has been correlated with smoking for Whites but not Blacks.[10–12] Correlates of regular smoking also differed for Black and White samples suggesting that the causes of smoking for these groups differ as well.[13]

This study examines whether the correlates and therefore possible causes of adolescent smoking are the same for Blacks and Whites. The correlates of smoking initiation are first examined for combined samples, testing for the effects of race, then examined separately for Black and White samples.

Methods

Sampling and Survey Procedures

The data are from a field experiment conducted to evaluate the influence of a mass media campaign on the onset of adolescent cigarette smoking. The campaign and its effectiveness are described elsewhere.[14,15] Ten of the 81 Standard Metropolitan Statistical Areas (SMSAs) in the southeastern United States that were similar on social and demographic characteristics and located in nonoverlapping broadcast areas were selected for the experiment. Using cluster sampling procedures, approximately 80 clusters of households were randomly selected in each SMSA to yield an expected average of 224 eligible adolescents per SMSA. Baseline surveys, conducted from April 1 through October 13, 1985, included all eligible adolescents (ages 12–14) within each SMSA (N = 2,534). In the presence of a trained interviewer, adolescents completed self-administered questionnaires in their homes. Follow-up data were collected two years later. For these analyses, data are aggregated for the 10 SMSAs.

Measurement

Cigarette smoking was derived from a 13-category item that ranged from "I have never smoked" to "I smoke two or more packs per day." Adolescents who smoked at least "a few times in the past year" were considered smokers and adolescents who had not smoked in the past year including never smokers were considered nonsmokers. A binary variable rather than the full range was used because the distribution was skewed toward infrequent and light smoking. This defini-

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From the Department of Health Behavior and Health Education (Headen, Bauman), the Department of Sociology and Carolina Population Center (Deane), and the Department of Biostatistics (Koch), all at the University of North Carolina, Chapel Hill.

Address reprint requests to Sandra W. Headen, PhD, Department of Health Behavior and Health Education, School of Public Health, University of North Carolina, Chapel Hill, NC 27599-7400. This paper, submitted to the Journal March 29, 1990, was revised and accepted for publication January 8, 1991.

tion of smoking is consistent with studies that focus on the first or second smoking experience among young adolescents who are experimenting with cigarettes.[4,5] To enhance the validity of self-reports of cigarette smoking, adolescents who completed questionnaires were told that they would be asked to provide aveolar breath and saliva samples to measure actual smoking behavior.[18,19]

Explanatory variables that had been identified as correlates of adolescent smoking in prior research were selected for study.[1,2,4,5,13,18-22] These variables are associated with the social learning and problem behavior theoretical perspectives and are described in the appendix.[7,10,11,21] Age, gender, and parent education were included as control variables. The following transformations were performed:

- Parental education, age, educational intentions, and availability of cigarettes were included as binary variables because the distribution of responses for their original categories were skewed.
- Religiosity, peer versus parental influence, parental permissiveness, and rebelliousness were included as continuous scales and derived by obtaining the means of several items. Closeness to mother and mother's approval of adolescent smoking were included in their original forms.

Analysis

To examine smoking initiation directly, only the 1,277 subjects from the 1985–87 panel who had not smoked within the year preceding 1985 were included in these analyses, eliminating the possible confounding effects of prior smoking behavior on 1987 measures.

Evaluation of correlates of adolescent smoking was primarily based on a logistic regression analysis that included race and 10 other explanatory variables as predictors of smoking initiation. This model was a simplification of another model that had included 15 explanatory variables. Four variables—intentions to smoke, parental permissiveness, mother's approval of adolescent smoking, and availability of cigarettes—were dropped from the earlier model. The −2 Log likelihood statistic showed that the explanatory capacity of the models that included and excluded these variables was similar. The number of subjects who indicated an intention to smoke at baseline was too small for reliable estimation of its effect. The remaining variables were included to examine whether rates of maternal smoking reported by Black adolescents may be associated with social environments characterized by parental leniency, easy access to cigarettes, and tolerance of adolescent smoking. The effects of these variables were expected to be relatively small and essentially no explanatory ability for them was confirmed after adjustment for other variables in the model.

This study reports findings of logistic regression analyses for Black and White samples that resulted from the following procedures. For the combined samples, stepwise procedures were used to identify significant two-way interactions among 11 explanatory variables. Significant interactions with race were of particular interest because of their implication of different relationships for Blacks and Whites. In a subsequent step, three-way interactions, including race and explanatory variables involved in significant two-way interactions, were also examined. Finally, the model that had been obtained on the combined samples was evaluated for the separate samples of Black and White adolescents. The significant interactions between race and the variables, friend smoking and age, in the equation for the combined samples provided the justification for these analyses.

Results

Subject Participation

Of the 2,534 adolescents who were eligible for study, 2,102 participated in 1985 and 1,637 participated in both 1985 and 1987. Among 429 subjects who were eligible but did not participate in 1985, 50.7 percent (n = 218) refused and 49.3 percent (n = 211) had moved from the area, could not be reached after multiple attempts, or were excluded for other reasons. Failure to locate subjects and missing information were the primary reasons for attrition by 1987. Other analyses of these data showed that subjects in the panel and those who failed to complete the follow-up questionnaire did not differ on the following dimensions: age, gender, educational intentions, religiosity, rebelliousness, participation in school activities, and parental closeness.[23] However, subjects lost to follow-up were more likely to be regular smokers at baseline than adolescents who participated at both data collection points (5.6 percent versus 3.1 percent, respectively).[15]

TABLE 1—1987 Smoking Status of 1985 Nonsmokers

Characteristics	Total N	Smoker Percent	Nonsmoker Percent
12 years old			
Female			
Black	63	9.5	90.5
White	130	23.1	76.9
Male			
Black	56	7.1	92.9
White	143	26.6	73.4
Subtotal	392	19.9	80.1
13 years old			
Female			
Black	58	13.8	86.2
White	164	13.4	86.6
Male			
Black	84	11.9	88.1
White	135	24.4	75.6
Subtotal	441	16.6	83.4
14 years old			
Female			
Black	63	20.6	79.4
White	150	26.7	73.3
Male			
Black	67	17.9	82.1
White	142	30.3	69.7
Subtotal	422	25.6	74.4
Total		20.6	79.4
Black	391		
White	864		
Grand Total	1,255		

TABLE 2—Descriptive Statistics for Explanatory Variables (1985)

Variables	Means or Proportions		Difference (95% CI)
	Black	White	
Best friend smokes[a]	22.90%	26.90%	4.00 (3.94, 4.06)
Mother smokes[a]	40.70%	27.30%	13.35 (13.32, 13.38)
Religiosity[b]	2.40	2.30	.10 (.02, .18)
Rebelliousness[c]	.22	.23	.01 (−.05, .07)
Close to mother[d]	4.58	4.48	.10 (00, .20)
Peer/parent influence[e]	3.30	2.60	.70 (.58, .82)
Educational intentions[f]			
college impossible[g]	8.20%	2.20%	6.00 (5.96, 6.04)
college possible[h]	34.80%	24.50%	10.20 (10.10, 10.30)

Categories were:
a) 1 = smoker, 0 = nonsmoker;
b) 3 = religion important, 0 = religion unimportant;
c) 1 = rebellious, 0 = nonrebellious;
d) 5 = extremely close to mother, 1 = not at all close;
e) 4 = parents more important, 0 = peers more important;
f) distributions for reference category (college certain) were 57% and 73.3% for Blacks and Whites respectively;
g) 1 = college impossible, 0 = college certain;
h) 1 = college possible, 0 = college certain.

TABLE 3—Adjusted Odds Ratios (and 95% confidence intervals) for Logistic Regression of Explanatory Variables on Smoking Initiation

Variables	Adjusted Odds Ratios (95% CI)	
	Blacks (n = 401)	Whites (n = 876)
Best friend smokes[a]	1.05 (.53, 2.10)	2.44 (1.68, 3.53)
Mother smokes[a]	.01 (.00, .76)	.06 (.01, .59)
Religiosity	.83 (.56, 1.22)	.81 (.65, 1.01)
Rebelliousness	1.34 (.37, 4.85)	2.39 (1.17, 4.85)
Close to mother	.60 (.39, .92)	.76 (.59, .98)
Peer/parent influence	1.04 (.75, 1.45)	.80 (.68, .94)
Educational intentions[b]		
College impossible	2.10 (.79, 5.58)	1.73 (.60, 5.00)
College possible	1.25 (.63, 2.48)	1.40 (.95, 2.08)
Parent education[c]		
< High school	.87 (.31, 2.46)	.30 (.12, .75)
College graduate	1.27 (.54, 3.00)	.81 (.55, 1.20)
Age[d]		
13 years old	1.12 (.40, 3.10)	.51 (.31, .84)
14 years old	2.53 (1.14, 5.64)	.93 (.62, 1.40)
Gender[e]	1.17 (.58, 2.39)	.59 (.41, .85)
Race[f]		

High scores mean more of a quality or parents influential. Reference categories were:
a) nonsmoker,
b) college certain,
c) high school graduate,
d) 12 years old,
e) male,
f) white.
Model Chi square = 25.74, df = 16, ($p < .06$), and 100.43, df = 16, ($p < .001$) for Black and White samples respectively.

Racial Differences in Smoking Initiation

Associations among race, gender, and 1987 smoking status of subjects who were nonsmokers in 1985 were examined separately for 12, 13, and 14 year olds. The results, shown in Table 1, indicate that, with the exception of 13 year old females, more White than Black adolescents initiated smoking during this two-year period. Smoking initiation more than doubled with age among Black adolescents, but increased only slightly with age among White adolescents.

Black-White Differences in Explanatory Variables

Results shown in Table 2 reveal some notable differences between Black and White adolescents on the explanatory variables. More Blacks than Whites reported that their mother smokes, but fewer Blacks reported having a best friend who smokes. Also, Black adolescents considered parents more important influences than did White adolescents. The majority of adolescents were confident that college was in their future, although fewer Blacks then Whites thought so. More Blacks than Whites were tentative about college attendance (college possible) or reported that a college education would be impossible for them.

Correlates of Smoking Initiation

The results of logistic regression analyses for Black and White samples reveal the following main effects, shown in Table 3, and several interactions that are discussed later:

- Friend behavior was more strongly related to smoking initiation among White than among Black adolescents; the odds of smoking if a friend smoked were 2.44 times greater among White adolescents but did not increase among Black adolescents.
- The age of smoking initiation also differed for Blacks and Whites; compared with Blacks, Whites were more likely to start smoking at age 12; Blacks were more likely to start smoking at age 14 than at ages 12 or 13.

Significant interactions in these equations revealed several findings. For both Black and White samples the odds of initiating smoking if mother smoked were extremely small among adolescents who reported being distant from mother (.03 and .10 for Blacks and Whites, respectively) and approached unity among adolescents who reported being close to mother (1.02 and .98 for Blacks and Whites, respectively). Another finding was the difference in smoking initiation among White males and females from low versus higher socioeconomic backgrounds. The interaction between gender and parent education in the equation for the White sample indicated that when neither parent graduated high school, males were 3.53 times as likely as females to initiate smoking; when one parent completed high school or college, males were only .58 times as likely to initiate smoking compared with females.

Discussion

The results of this study reveal two important findings regarding racial differences in adolescent smoking initiation. The first is that friend smoking is related to smoking initiation among Whites but not among Blacks. This observation suggests that smoking prevention programs that focus on the attitudes and behaviors of peers may not be effective with Black adolescents.

The second important finding is that Whites initiated smoking earlier than Blacks, an observation that is consistent with results of other studies conducted in the southeastern United States.[10,11,18] This outcome suggests that smoking prevention programs should be implemented in elementary grades to deter smoking among Whites and should continue into middle school to prevent smoking initiation among Blacks.

The results of this study suggest that racial differences in adolescent smoking can be examined more effectively with stratified analyses; aggregating data for Blacks and Whites obscures differences between these groups. For example, equations for both Black and White samples showed that adolescents who felt close to their mothers did not initiate smoking if their mother smoked. Also, the differences in smoking initiation among males and females from low versus higher socioeconomic backgrounds were found for the White sample only.

Our results also show that smoking rates and sample sizes must be sufficiently large to detect meaningful relationships for the Black sample when stratified analysis is used. The small proportion of smokers among Black adolescents in this study (14 percent of Blacks smoked in 1987 compared with 25 percent of Whites) and the smaller size of the Black sample decreased the capacity of logistic regression to explain smoking among Blacks.

In summary, this study found that Black adolescents initiated smoking later than Whites and were not influenced by their friends to start smoking. These findings indicate the need for more research on how differences in correlates of adolescent smoking contribute to differences in smoking among Blacks and Whites. We suggest greater emphasis on identifying and testing models of adolescent smoking that consider distinctive cultural and environmental factors that shape the behavior of Black adolescents. □

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APPENDIX

Explanatory Variables

Variable names and categories	Description and original form
Best friend smokes Mother smokes 0 = nonsmoker (category 1) 1 = nonsmoker (categories 2–5)	Cigarettes smoked per day; five-point scale from 1 = none to 5 = more than one pack per day
Intentions to smoke 0 = do not intend to smoke (categories 1–3) 1 = intend to smoke (categories 4, 5)	Probability of smoking in two years; five-point scale from 1 = sure I will not smoke to 5 = sure I will smoke
Availability of cigarettes 0 = hard to get (categories 1, 2) 1 = easy to get (categories 3, 4)	Ease of getting cigarettes; four-point scale from 1 = very hard to 4 = very easy
Mother's approval of adolescent smoking	How mother would feel if adolescent smoked; five-point scale from 1 = strongly disapprove to 5 = strongly approve
Educational intentions 1 0 = college certain (categories 4, 5) 1 = college impossible (category 1)	How likely is college? five-point scale from 1 = attendance, impossible to 5 = attendance certain
Educational intentions 2 0 = college certain (categories 4, 5) 1 = college possible (categories 2, 3)	(same as above)
Religiosity; four-point scale from 0 = religion not important, never go to services to 3 = religion important, services weekly	Derived from mean of two, four-point scales: religion important? 1 = not important to 4 = very important; how often services attended? 1 = never to 4 = once a week or more
Rebelliousness continuous variable; range 0, 1 0 = nonrebellious 1 = rebellious	Derived from mean of four binary items; 1 = yes, 2 = no; try to follow rules? break rules you disagree with? break rules for sake of it? break rule to prove a point?
Close to mother; five-point scale from 1 = not close to 5 = extremely close	Feel close to mother?
Peer/parent influence; five-point scale from 0 = friends more important to 4 = parents more important	Derived from two, five-point scales; who do you listen to when you have problems? who influences your behavior? from 1 = friends more often to 5 = parents more often (3 = friends/parents equally)
Parental permissiveness/independence; continuous scale, range 0, 1	Derived from mean of five binary items; 1 = yes, 2 = no; weekend curfew? choose own friends? choose own clothes? regular bedtime? say parents are strict?
Parent education 1 0 = high school graduate (categories 2, 3, 4) 1 = less than high school (category 1)	Highest grade of either parent; six-point scale from 1 = did not graduate from high school to 6 = more than four years of college
Parent education 2 0 = high school graduate (categories 2, 3, 4) 1 = college graduate or higher (categories 5, 6)	(same as above)

Health Behaviors, Risk Factors, and Health Indicators Associated with Cigarette Use in Mexican Americans: Results from the Hispanic HANES

David J. Lee, PhD, and Kyriakos S. Markides, PhD

ABSTRACT

Background: Cigarette smokers often engage in other, potentially deleterious, health behaviors. Such behaviors have not been well documented in Mexican American smokers.

Methods: Data from the Southwestern sample of the Hispanic Health and Nutrition Examination Survey (HHANES) were employed to investigate differences in health behaviors, risk factors and health indicators between cigarette smokers and nonsmokers among Mexican Americans. Differences between those smoking < 10 and 10 or more cigarettes per day were also examined by age group and gender.

Results: Positive associations between smoking status and heavy coffee and alcohol consumption were found across gender and age groups. Less consistent was the finding that smokers weighed less than nonsmokers. Lower systolic and diastolic blood pressures in middle-aged smokers, and higher levels of depressive symptomatology among smoking women were found. Those smoking 10 or more cigarettes per day were more likely to report heavy coffee consumption, with younger men reporting greater activity limitation due to poor health. Middle-aged men and women in the 10+ category were generally in better health than lighter smokers.

Conclusions: Modest associations between cigarette smoking, health behaviors and risk factors found in other studies were confirmed in this Mexican American population. Few significant associations between smoking and health status were noted. (*Am J Public Health.* 1991;81:859–864)

Introduction

Cigarette smoking is the leading cause of preventable mortality in the United States with approximately 390,000 deaths annually attributed to this health behavior.[1] The prevalence of smoking varies by age, gender, and ethnicity. Males have higher rates of smoking than do women although the gap in rates has narrowed in recent years.[2] Smoking rates climb dramatically in adolescence, peak in middle age, and decline in old age. Overall, smoking rates among Mexican American men are comparable to estimates for Black American males (44 versus 41 percent)[3,4] while rates for Mexican American women are comparable to Caucasian women (25 versus 27 percent).[3,4] However, the prevalence of heavy cigarette consumption is lower among both male and female Mexican Americans than in other ethnic groups.[5]

While much research has focused on basic epidemiologic relationships between smoking and health outcomes, some research has focused on the relation between smoking and other risk factors. The importance of this research is four-fold:

- smoking may interact with other risk factors producing synergistic increases in the risk of negative health outcomes[6];
- smokers attempting to quit often need to reduce exposure to other health behaviors (e.g. alcohol consumption) which prompt smoking behavior;
- smokers who engage in multiple negative health behaviors are less likely to successfully quit smoking[7];
- change in one health behavior may lead to concomitant reductions in other health behaviors.[8,9]

Much of this research has focused on interrelations between alcohol, cigarette, and caffeine use documenting modest associations among these psychoactive substances.[10] Cigarette consumption has also been found to positively correlate with cholesterol levels,[11] but paradoxically is associated with lower blood pressure in adults.[12,13] There is also an inverse relationship between smoking and obesity.[14]

Although cigarette smoking is correlated with a number of deleterious health outcomes,[15] both gender and aging differences can influence these associations.[16] Therefore it is important to examine these differences in different age and gender groups. In addition, because most individuals do not start smoking after the age of 25, associations between smoking status and health may be influenced by two factors: those who quit smoking due to poor health; and selective survival with those most vulnerable to smoking-related mortality dying during middle age.[17] Cross-sectional analyses often reveal that older smokers are in better health than nonsmokers.[18]

The purpose of this study is to examine health behavior, risk factor, and health indicator correlates of smoking in a representative sample of Mexican Americans. Because of important age and gender differences, these associations will be examined separately by gender and age groups.

Address reprint requests to David J. Lee, PhD Department of Epidemiology and Public Health, School of Medicine, University of Miami, P.O. Box 016069 (R-669), Miami, FL 33101. Dr. Markides is with the Department of Preventive Medicine & Community Health, University of Texas Medical Branch, Galveston. This paper, submitted to the Journal July 24, 1990, was revised and accepted for publication January 23, 1991.

Methods

The data employed are from the Hispanic Health and Nutrition Examination Survey (HHANES) conducted during 1982–84 by the National Center for Health Statistics.[19] Our analysis is limited to Mexican American respondents from the southwestern states (Texas, New Mexico, Colorado, Arizona, and California). The HHANES employed a complex multistage sampling design that aimed at obtaining a representative sample of Mexican Americans in the southwest (as well as Cuban Americans in the Miami (Florida) area and Puerto Rican Americans in the New York City area) ages 6 months to 74 years. The sampling design employed assured a 97 percent representation of the Mexican origin population of the southwest as well as 84 percent of all persons of Mexican origin in the United States (for details on sampling procedures see Gonzalez).[20]

Our analysis is limited to persons ages 20 to 74. A total of 3,935 persons in this age group participated in the household interview, with 3,326 participating in the physical examination where data on body weight and height, blood pressure, cholesterol, alcohol consumption, and depression were collected.

Measures

Current smoking status was ascertained in all participants who indicated that they had smoked at least 100 cigarettes in their lifetime. Current smokers were asked how many cigarettes they smoked daily. Responses ranged from less than one to greater than 86 with the actual number recorded between these two extremes. The median number of cigarettes consumed in this sample was used to compare participants whose consumption was "moderate" (<10 cigarettes per day) and "immoderate" (≥10 cigarettes per day).

Measures of health behaviors and risk factors included *Heavy Coffee Consumers* (coded 1 for those consuming 4 or more cups per day, 0 for less than 4, collected during a three-month dietary recall interview); *Current Alcohol Consumer* (coded 1 for current alcohol consumer; 0 for abstainer); *Average Alcohol Consumption* (average number of drinks consumed weekly); *Body Mass Index* [weight (in kilograms) divided by the square of the height of the individual (in meters)[21]]; *Systolic* and *Diastolic Blood Pressure* (based on the average of two blood pressure assessments during the physicians's examination[22]) and serum *Cholesterol* (milligrams per deciliter of blood plasma).

Health indicators included in this analysis were Activity Limitation, Self-Rated Health, Physician's Assessment, and Depression. *Activity Limitation* was based on a series of questions regarding limitation in major activity (e.g. job, housework, etc.) because of an impairment or health problem. In addition to the above, persons 71 to 74 years of age were considered to have activity limitation if they needed the help of others with personal care needs or daily routine activities because of any impairment or health problem, (coded 1 for activity limitation, 0 for other). *Self-Rated Health* was based on a single item asking respondents to rate their health as excellent, very good, good, fair, or poor (coded 1 to 5). The *Physician's Assessment* also graded subjects on the same 1 to 5 scale for an overall assessment of health. *Depression* was measured by the Center for Epidemiologic Studies Depression (CES-D) Scale which has been widely used in community surveys and has been extensively validated,[23,24] including with Hispanic populations.[25,26] Internal consistency reliability for the scale with the current data was in excess of .85. Range of scores on the CES-D is 0–60 with higher scores indicative of greater symptomatology.

Analysis

All analyses were calculated separately for males and females in three age groups: 20–39; 40–59; 60–74 years. Due to small sample sizes examination of the smoking consumption groups (<10; 10+) were not undertaken in the 60–74 age group. Log linear and general linear modeling were used to compare mean and percent differences across smoking categories.[27] When age was significantly different across smoking status or level of consumption within these age and sex categories, age-adjusted means and percents were calculated. For continuously distributed variables this analytic approach paralleled that of multiple classification covariance analysis[28] while for dichotomous dependent variables it parallels binary multiple regression analysis.[29] Both approaches require that the covariant be homogeneous across categories of the independent variable. Significant interactions between the covariate and the factor indicate that age adjustment may produce misleading results. Therefore, interactions between age and smoking status and level of consumption were first run to confirm independence. (See Mendes de Leon and Markides[26] for further details on application of this approach).

Analyses were performed with adjustments for sample weights and design effects. The latter were necessary because the HHANES employed a complex multistage sampling design. Sample weights were normalized by multiplying the number of subjects in each group employed in the analysis with each respondent's weight divided by the sum of the weights for the group. Design effects were estimated using the Taylor Approximation method of the SESUDAAN program and were used to adjust variances in the analyses.[30]

Results

The prevalence of smoking and level of cigarette consumption (<10; 10+) by age group and gender is reported in Table 1. Males were more likely to be smokers ($\chi 2 = 89.17$, $p<.001$) and to smoke more cigarettes ($\chi^2 = 7.93$, $p<.01$) than females. Prevalence rates also varied by age group with the highest rates found in the middle aged group for both smoking status ($\chi^2 = 7.11$, $p<.05$) and amount of cigarette consumption ($\chi^2 = 9.11$, $p<.05$). The age group by gender status interactions were not significant indicating that the smokers did not disproportionately fall into specific age group and gender categories.

Table 2 presents frequencies and means for health behaviors, risk factors, and health indicators for young males (ages 20–39). Smokers were more likely to report current alcohol consumption, reported greater consumption of alcohol, were more likely to be heavy coffee consumers, and reported poorer self-rated health. Smokers consuming 10 or more cigarettes per day were more likely to be heavy coffee consumers and reported greater activity limitation than smokers consuming less than 10 cigarettes per day.

Table 3 presents frequencies and means for health behaviors, risk factors, and health indicators for young females. As with young males, female smokers were more likely to report current alcohol consumption and were heavier drinkers. Levels of depressive symptomatology were also significantly higher among smokers than non-smokers.

Females consuming 10 or more cigarettes per day, were significantly older than smokers consuming less than 10 cigarettes per day (29.82 versus 28.22). Therefore, after confirming no age by consumption status interactions, age-adjusted means were calculated. As can be seen in

TABLE 1—Prevalence of Smoking Status and Prevalence of Smokers Consuming 10 or More Cigarettes Per Day by Gender and Age Group

		Percent Smokers			Percent Smoking ≥10 Cigarettes/Day	
	N	Male	Female	N	Male	Female
Age Groups						
20–39	2163	42.5	24.2	704	51.5	46.6
40–60	1313	46.4	27.4	455	60.9	50.3
60–74	457	39.7	19.8	133	74.1	54.0

Note: Main effects for smoking status: Age: $\chi^2 = 7.11$; $p<.05$; Gender: $\chi^2 = 89.17$; $p<.001$; Interaction: $\chi^2 = .11$; p = n.s
Main effects for cigarette consumption: Age: $\chi^2 = 9.11$; $p<.05$; Gender: $\chi^2 = 7.93$, $p<.01$; Interaction: $\chi^2 = 2.02$, p = n.s

TABLE 2—Comparison of Nonsmokers and Smokers and Amount of Cigarette Consumption Across Health Behaviors, Risk Factors, and Health Indicators Among Young Men (ages 20–39)

	Smoking Status		Amount of Smoking	
	Nonsmokers (N = 445–586)	Smokers (N = 319–436)	<10/Day (N = 148–212)	≥10/Day (N = 170–221)
	Mean	Mean	Mean	Mean
Health Behaviors/Risk Factors				
% Heavy Coffee Consumers	3	6*	1	10***
% Current Alcohol Consumers	68	79**	81	78
Average Alcohol Consumption^ (Drinks/Week)	4.4	5.5**	5.4	5.7
Body Mass Index (kg/m^2)	26.0	25.3	25.4	25.2
Systolic Blood Pressure (mmHg)	117.6	118.3	118.6	117.8
Diastolic Blood Pressure (mmHg)	76.3	75.8	75.6	76.0
Cholesterol (mg/dl)	194.9	197.2	193.3	200.4
Health Indicators				
% with Activity Limitation	9	9	5	13**
Self-Related Health	2.6	2.9***	3.0	2.8
Physician's Assessment	1.4	1.4	1.4	1.4
CES-D Depression	6.3	7.1	6.8	7.3

Note: *$p<.05$; **$p<.01$; ***$p<.001$
^Based on current consumers only (N = 319 nonsmokers; N = 264 smokers; N = 126 <10 cigarettes/per day; N = 137 10+ cigarettes/per day smokers).

Table 3, there were no significant health behavior, risk factor, or health indicator differences between individuals in the two smoking categories.

Table 4 shows the middle-aged (40–59) male smokers were more likely to report greater levels of alcohol consumption, weighed less, reported poorer self-rated health, and had lower systolic and diastolic blood pressure. Smokers consuming 10 or more cigarettes per day had significantly lower diastolic blood pressure than smokers consuming less than 10 cigarettes per day.

Table 5 presents frequencies and means for health behaviors, risk factors, and health indicators for middle-aged females. Age adjustment of means was necessary for the smoking status analyses after determining that the mean age of the smoking group was significantly higher than the nonsmoking group (49.20 versus 48.00). There were no significant interactions between smoking status and age.

Smokers were more likely to be current alcohol consumers, heavy coffee consumers, weighed less, and had lower diastolic blood pressure. Smokers also reported greater depressive symptomatology than nonsmokers. Smokers consuming 10 or more cigarettes per day weighed less, reported better self-rated health and less activity limitation. Table 6 shows that older (60–74) male smokers were more likely to report current alcohol consumption, were more likely to be heavy coffee consumers, and weighed less. Older female smokers were more likely to report current alcohol consumption and were more likely to be heavy coffee consumers.

Discussion

The most consistent differences between smokers and nonsmokers and between those in the two smoking categories were with alcohol and coffee consumption—findings which are in agreement with studies in other populations.[10] In general, smoking status was negatively associated with BMI, a finding also consistent with previous research.[14] Blood pressure was generally lower in smokers than in nonsmokers, although this association was significant only for middle-aged smoking men and women. More consistent associations between smoking status as well as the presence of a dose-response relationship have been reported in other studies.[12]

No significant differences were found for cholesterol levels across smoking groups. The average reported difference in cholesterol levels in smokers and nonsmokers in other studies is approximately 3 percent.[11] In the present analysis, this difference was exceeded only in comparisons of those consuming 10 or more cigarettes per day with those smoking less than this amount in middle-aged women and young men.

Findings relating to physical and emotional health were equivocal, varying by age group and gender. There was limited support that health was poorer in younger, but not older males. In prospective analyses, Hirdes, *et al.*[18] found that smoking status was associated with future risk of disability although this association failed to reach statistical significance. Lack of a strong relationship between morbidity and cigarette consumption in this and other studies may reflect selective survival with the heaviest most vulnerable smokers dropping out of the cohort due to premature mortality.[17]

Our findings have implications for epidemiologic investigations of smoking. Associations between smoking and other health behaviors reinforce the need to control for the effect of these behaviors when examining the impact of smoking on health. Psychological status may also be important to control for in such epidemiologic investigations. Depression, found to be associated with smoking status among women in this study and in previous research,[31,32] is also a risk factor for mortality[33]—particularly among cigarette smokers.[34] Researchers may consider controlling for depression in investigations of smoking and health outcomes as well as exploring possible interactions between smoking and depression.[34,35]

Limitations of this study include potential nonresponse bias and reliance on self-report measures of cigarette consumption. Eighty-seven percent of Mexican Americans approached to participate in the HHANES completed the household interview while 76 percent completed the physician's examination.[36] It should be noted, however, that the weighting procedure incorporated in the present analysis includes adjustment for nonresponse.[36]

Although national estimates of self-reported cigarette consumption are correlated with cigarette sales in the United States (average r = .72),[37] specific analysis of self-report data from the HHANES suggest that underreporting of consumption may have occurred. Based on serum levels of cotinine, a principle metabolite of nicotine, Perez-Stable, *et al*,[38] estimated that 20–25 percent of those reporting 1–9 cigarettes used per day were actually underreporting cigarette consumption, while underreporting estimates for those reporting greater than 10 cigarettes per day ranged from 2–11 percent. These findings suggest that a proportion of subjects in the 1–9 cigarette consumption category in the present analysis may have been misclassified due to underreporting.

In general, this paper confirmed the modest associations found between cigarette smoking and health behaviors and risk factors found in other populations, although these associations did vary by age and gender. Furthermore, few significant associations between smoking and health

TABLE 3—Comparison of Nonsmokers and Smokers and Amount of Cigarette Consumption Across Health Behaviors, Risk Factors, and Health Indicators Among Young Women (ages 20–39)

	Smoking Status		Amount of Smoking⁺⁺	
	Nonsmokers (N = 733–868)	Smokers (N = 229–271)	<10/Day (N = 129–147)	≥10/Day (N = 100–124)
	Mean	Mean	Mean	Mean
Health Behaviors/Risk Factors				
% Heavy Coffee Consumers	2	4	3	6
% Current Alcohol Consumers	28	54***	52	56
Average Alcohol Consumption^ (Drinks/Week)	2.8	3.9***	3.6	4.2
Body Mass Index (kg/m^2)	25.9	25.9	26.3	25.4
Systolic Blood Pressure (mmHg)	108.8	109.2	109.1	109.5
Diastolic Blood Pressure (mmHg)	70.6	70.1	70.1	70.3
Cholesterol (mg/dl)	190.9	189.0	188.8	190.0
Health Indicators				
% with Activity Limitation	11	12	15	10
Self-Related Health	2.9	2.9	2.8	3.0
Physician's Assessment	1.4	1.5	1.5	1.5
CES-D Depression	8.8	10.7*	10.8	10.5

Note: *p<.05; ***p<.001

^Based on current consumers only (N = 205 nonsmokers; N = 124 smokers; N = 67 <10 cigarettes/per day; N = 57 10+ cigarettes/per day. ⁺⁺Age-adjusted means are reported.

TABLE 4—Comparison of Nonsmokers and Smokers and Amount of Cigarette Consumption Across Health Behaviors, Risk Factors, and Health Indicators Among Middle-Aged Men (ages 40–59)

	Smoking Status		Amount of Smoking	
	Nonsmokers (N = 247–304)	Smokers (N = 201–265)	<10/Day (N = 81–103)	≥10/Day (N = 118–159)
	Mean	Mean	Mean	Mean
Health Behaviors/Risk Factors				
% Heavy Coffee Consumers	5	9	5	12
% Current Alcohol Consumers	67	75	80	72
Average Alcohol Consumption^ (Drinks/Week)	3.8	5.1**	5.6	4.7
Body Mass Index (kg/m^2)	27.7	26.0***	26.1	26.0
Systolic Blood Pressure (mmHg)	127.1	123.0*	125.7	121.5
Diastolic Blood Pressure (mmHg)	83.3	80.0**	82.6	78.4**
Cholesterol (mg/dl)	221.2	219.1	220.1	218.4
Health Indicators				
% with Activity Limitation	22	22	21	23
Self-Related Health	3.0	3.2*	3.2	3.1
Physician's Assessment	1.8	1.8	1.8	1.8
CES-D Depression	5.0	6.2	7.2	5.5

Note: *p <.05; **p<.01; ***p<.001

^Based on current consumers only (N = 163 nonsmokers; N = 157 smokers; N = 68 <10 cigarettes/per day; N = 89 10+ cigarettes/per day.

TABLE 5—Comparison of Nonsmokers and Smokers and Amount of Cigarette Consumption Across Health Behaviors, Risk Factors, and Health Indicators Among Middle-Aged Women (ages 40–59)

	Smoking Status^{++}		Amount of Smoking	
	Nonsmokers (N = 447–546)	Smokers (N = 162–196)	<10/Day (N = 79–96)	≥10/Day (N = 79–97)
	Mean	Mean	Mean	Mean
Health Behaviors/Risk Factors				
% Heavy Coffee Consumers	5	9*	6	14
% Current Alcohol Consumers	21	44*	39	54
Average Alcohol Consumption^ (Drinks/Week)	2.4	3.7	3.0	4.4
Body Mass Index (kg/m^2)	28.7	27.5*	29.0	25.9***
Systolic Blood Pressure (mmHg)	123.0	120.7	119.7	119.7
Diastolic Blood Pressure (mmHg)	77.6	74.8***	75.6	73.9
Cholesterol (mg/dl)	216.4	222.1	215.0	225.8
Health Indicators				
% with Activity Limitation	23	22	26	15*
Self-Related Health	3.3	3.3	3.4	3.0*
Physician's Assessment	1.9	1.8	1.8	1.8
CES-D Depression	8.2	12.8***	12.8	12.4

Note: *p<.05; **p<.01; ***p<.001

^Based on current consumers only (N = 96 nonsmokers; N = 74 smokers; N = 29 <10 cigarettes/per day; N = 44 10+ cigarettes/per day. $^{++}$Age-adjusted means are reported.

TABLE 6—Comparison of Nonsmokers and Smokers Across Health Behaviors, Risk Factors, and Health Indicators Among Older Men and Women (ages 60–74)$^{++}$

	Smoking Status: Men		Smoking Status: Women	
	Nonsmokers (N = 89–121)	Smokers (N = 62–82)	Nonsmokers (N = 145–202)	Smokers (N = 37–51)
	Mean	Mean	Mean	Mean
Health Behaviors/Risk Factors				
% Heavy Coffee Consumers	0	9***	0	9***
% Current Alcohol Consumers	43	63*	11	30**
Average Alcohol Consumption^ (Drinks/Week)	2.9	4.5	—	—
Body Mass Index (kg/m^2)	26.9	25.5*	28.4	27.3
Systolic Blood Pressure (mmHg)	140.1	135.8	138.5	133.9
Diastolic Blood Pressure (mmHg)	80.2	78.2	77.6	74.6
Cholesterol (mg/dl)	214.2	213.6	234.1	235.2
Health Indicators				
% with Activity Limitation	42	30	37	41
Self-Related Health	3.5	3.4	3.7	3.5
Physician's Assessment	2.1	2.1	2.2	2.1
CES-D Depression	5.6	6.1	8.9	12.0

Note: *p<.05; **p<.01; ***p<.001

^Based on current consumers only (N = 42 nonsmokers; N = 40 smokers).

$^{++}$Examination of differences in amount of cigarette consumption was not undertaken due to small sample size.

status were found. Associations between smoking status and depression were found for women, but not men. These findings suggest that smoking researchers examining Mexican American populations also need to control for relevant risk factors and health behaviors associated with smoking when examining relationships between smoking and health. ☐

Acknowledgments

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Standardized Comparisons of the Use of Alcohol, Drugs, and Cigarettes Among Military Personnel and Civilians

Robert M. Bray, PhD, Mary Ellen Marsden, PhD, and Michael R. Peterson, DrPH

ABSTRACT

Background: Understanding the effectiveness of policies and programs aimed at combating substance abuse in the military requires comparison with the civilian population from which military personnel are drawn.

Methods: Standardized comparisons of the use of alcohol, drugs, and cigarettes among military personnel and civilians were conducted with data from the 1985 Worldwide Survey of Alcohol and Nonmedical Drug Use among Military Personnel and the 1985 National Household Survey on Drug Abuse. The two data sets were equated for age and geographic location of respondents, and civilian substance use rates were standardized to reflect the sociodemographic distribution of the military.

Results: Military personnel were significantly less likely than civilians to use drugs, but were significantly more likely to use alcohol and cigarettes and to engage in heavy use of alcohol and cigarettes. Heavy drinking was especially likely among young military men. Military women were similar to military men in their smoking and drug use patterns.

Conclusions: Findings suggest that military policies and programs have been notably effective in reducing drug use, but that efforts to limit alcohol and cigarette use should be intensified. Military efforts directed against alcohol abuse should be targeted toward younger men, while smoking and drug prevention programs should be directed toward both men and women. (*Am J Public Health.* 1991;81:865–869)

Introduction

Alcohol misuse, nonmedical drug use, and cigarette use can have negative effects on the health, social life, family relationships, and work performance of users. Although these negative effects can occur within all segments of society, they may be especially serious for United States military personnel who have the unique mission to protect and defend the nation. Because military personnel are drawn from the civilian population, other things being equal, the prevalence and level of use of alcohol, drugs, and cigarettes of the two populations are expected to be similar. Certain differences between military and civilian life, however, may create substance use rates that differ. It has been argued that distinctive military conditions such as relocation overseas, separation from family, or a greater perceived acceptance of use may foster higher rates of use among military personnel than civilians.[1,2] Alternatively, military policies or programs directed toward the elimination of abuse, such as the adoption of a policy of zero tolerance toward drug use and the implementation of an extensive urinalysis screening program, may foster rates of use among military personnel that are substantially lower than among civilians.[3]

Direct comparisons of differences in rates of use among military personnel and civilians may not be meaningful because the sociodemographic compositions of the two populations differ. Military personnel are disproportionately young and male, characteristics generally associated with higher levels of use.[4] Unstandardized rates may not therefore provide adequate comparisons of the extent of differences in substance use between military personnel and civilians.

Several standardized comparisons of alcohol and drug use among military personnel and civilians have been conducted. Observed differences in rates of problem drinking between military personnel and civilians ages 21 to 59 in the late 1970s, for example, were found to be largely but not entirely accounted for by differences in the demographic compositions of military and civilian populations.[5,6] The prevalence of use of alcohol and drugs among military personnel and for civilians ages 18 to 25 was quite similar in 1980,[7] but in 1982 the prevalence of alcohol use was higher and the prevalence of marijuana and cocaine use was significantly lower for military than civilian men ages 18 to 25.[8]

These comparisons are somewhat limited in that they rest largely on measures of the prevalence of alcohol use that are inadequate to describe the quantity and frequency of use, and comparisons of drug use that are based only on the prevalence of use of specific drugs. Further, some comparisons have been based on civilian data that were collected at a different time than the military data,[7] and other studies compared populations that differed in important ways. Some studies, for example, considered both males and females while others considered only males; some considered only ages 18 to 25 while others con-

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From the Center for Social Research and Policy Analysis, Research Triangle Institute (Bray, Marsden) and the Office of Assistant Secretary of Defense/Health Affairs (Peterson).

Address reprint requests to Robert M. Bray, PhD, Center for Social Research and Policy Analysis, Research Triangle Institute, P.O. Box 12194, Research Triangle Park, NC 27709. This paper, submitted to the Journal September 26, 1989, was revised and accepted for publication January 8, 1991.

Editor's Note: See also related editorial p 837 this issue.

TABLE 1—Sociodemographic Characteristics of Military Personnel and Civilians, Aged 18–55

Characteristics	Comparison Population	
	Civilian (N = 4,894)	Military (N = 8,084)
Ages (years)		
18–25	25.9	49.3
26–34	28.7	32.3
35–55	45.4	18.4
Sex		
Male	49.8	90.6
Female	51.0	9.4
Race/Ethnicity		
White	79.8	72.6
Black	10.9	16.8
Hispanic	6.9	6.3
Other	2.5	4.3
Education		
Less than high school	17.8	1.3
High school grad/GED	38.6	44.6
Some college	21.6	36.0
College degree or beyond	22.0	18.1

Note: Estimates are unstandardized percentages. Ns (in parentheses) show the number of cases on which weighted data are based.
Civilian data source: 1985 National Household Survey on Drug Abuse.
Military data source: 1985 Worldwide Survey.

sidered a broader age range. No recent military/civilian comparisons of alcohol and drug use are available, and no standardized comparisons of cigarette use have been conducted. For these reasons, general conclusions about military/civilian differences cannot be drawn.

This paper reports results of standardized comparisons of alcohol, drug, and cigarette use among military personnel and civilians in 1985, the most recent date for which comparable data were available. Unlike earlier studies, findings are reported for a broad age range as well as for males and females.

Methods

Data Sources and Analysis Files

Data for military personnel were drawn from the 1985 Worldwide Survey of Alcohol and Nonmedical Drug Use among Military Personnel, sponsored by the US Department of Defense (DoD) and conducted by Research Triangle Institute.* Data on civilians were drawn from the 1985 National Household Survey on Drug Abuse (NHSDA), sponsored by the National Institute on Drug Abuse and conducted by the Institute for Survey Research at Temple University and the Department of Sociology at the University of Kentucky.[9] The two surveys differ in overall questionnaire content, order of item presentation, and field procedures. The 1985 Worldwide Survey, for example, was based on self-report questionnaires administered in group settings, while the 1985 NHSDA data were collected in face-to-face interviews. These differences suggest some caution in making comparisons between the two surveys, but both surveys were carefully conducted and are based on large probability samples.

The eligible population for the 1985 Worldwide Survey consisted of all US active-duty military personnel stationed across the world except recruits, service academy students, persons absent without leave, and persons who had a permanent change of station at the time of data collection. A probability sample was selected for the survey using a deeply stratified, two-phase design. Completed questionnaires were obtained from 17,328 active-duty military personnel for a response rate of 80 percent. Of the responses, 85 percent were obtained in group administrations of the survey at military installations; the remainder were obtained by mailing questionnaires to a subsample of nonrespondents from the group administrations. Estimates of the sociodemographic characteristics of the respondent population showed a close correspondence to sociodemographic characteristics of the actual DoD population values.

The target population for the 1985 NHSDA civilian survey consisted of the household population of the continental United States, ages 12 and older. The household population was sampled using a multistage area probability design with 100 primary sampling units. Individuals in group quarters (military installations, correctional institutions, college dormitories, and hospitals) and those who had no permanent residence (homeless people and residents of single rooms in hotels) were not included in the study. The survey was administered face-to-face, and 8,038 completed interviews were obtained for a response rate of 83 percent.

Before analyzing the data, adjustments to the two data sets were made for differences in residential location and age to ensure greater comparability. To equate for location, military data were limited to respondents residing in the continental United States. Active-duty military personnel living off base in civilian housing were excluded from the NHSDA. To equate for age, both data sets were limited to respondents who were ages 18 to 55 (99.8 percent of the military respondents were in this age range). The adjusted data file for military personnel consisted of 8,084 cases, and the adjusted data file for civilians was 4,894 cases.

Standardization Procedures

As shown in Table 1, the military population is younger, somewhat better educated, and more likely to be male or Black than the civilian population.

Direct standardization[10] was used for comparisons of military personnel and civilians. Civilian data were standardized to the joint distribution in the military of age (18 to 21, 22 to 25, 26 to 34, 35 to 55), race/ethnicity (White, Black, other), education (high school diplomas or less, more than high school), and sex. The standardized estimate for the civilian population is interpreted as the mean that would have been obtained if the civilian population had the same sociodemographic distribution as the military.

Both the Worldwide Survey and the NHSDA used complex, nested sampling designs. The estimates and the standard errors of the standardized civilian data and the military data were computed using

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*Survey methodology and survey findings for the 1985 Worldwide Survey are described in detail in Bray RM, Marsden ME, Guess LL, Wheeless SC, Pate DK, Dunteman GH, Iannacchione VG: 19[illegible]wide survey of alcohol and nonmedical drug use among military personnel. Report prepared for the Assistant Secretary of Defense (Health Affairs), by Research Triangle Institute, Research Triangle Park, NC, 1986.

TABLE 2—Standardized Comparisons of Alcohol Use in Past 30 Days Among Military Personnel and Civilians, by Age Group

	Age Group/Population					
	18–25 years		26–55 years		Total	
Group/Measure	Civilian	Military	Civilian	Military	Civilian	Military
Males	(N = 776)	(N = 1,645)	(N = 1,250)	(N = 5,885)	(N = 2,026)	(N = 7,530)
Any drinking	76.9	87.8***	78.7	81.6	77.8	84.6***
Heavy drinking	14.1	32.1***	9.6	12.9**	11.8	22.3***
Females	(N = 994)	(N = 210)	(N = 1,773)	(N = 334)	(N = 2,767)	(N = 554)
Any drinking	63.2	83.6***	63.9	73.7	63.5	79.0***
Heavy drinking	4.7	7.6	1.7	4.9	3.3	6.3
Total	(N = 1,770)	(N = 1,855)	(N = 3,023)	(N = 6,219)	(N = 4,793)	(N = 8,084)
Any drinking	75.5	87.3***	77.5	80.9*	76.5	84.1***
Heavy drinking	13.1	29.6***	8.9	12.2**	11.0	20.8***

Note: Estimates are percentages. Ns (in parentheses) show the number of cases on which weighted data are based. Civilian data have been standardized to the military data by age, education, race/ethnicity, and sex.

*Different from civilian at the .05 level.

**Different from civilian at the .01 level.

***Different from civilian at the .001 level.

Civilian data source: 1985 National Household Survey on Drug Abuse.

Military data source: 1985 Worldwide Survey.

RATIO2, the software package developed by Research Triangle Institute to analyze surveys with complex sampling designs.** T-tests were used to assess the significance of the differences between the standardized rate for the civilian population and the unstandardized rate for the military. Some of the substantial differences between military personnel and civilians have p values greater than 0.05 because of the relatively large standard errors.

Comparisons are reported for males, females, and both sexes combined and for those ages 18 to 25, 26 to 55, and the total group ages 18 to 55. Comparisons for the younger age group provide information about substance use among those at highest risk for use. For the military, this younger group consists largely of enlisted military personnel. The older age group is generally at lower risk, and in the military this group includes a higher proportion of officers.

Measures of Substance Use

The measures of alcohol use are the prevalence of any drinking and heavy drinking during the past 30 days. Any drinking refers to having had one or more drinks of beer, wine, or liquor during the past 30 days. Heavy drinking refers to consuming five or more drinks per typical drinking occasion at least once a week and is based on a drinking-level classification scheme adapted from Mulford and Miller.[11] The measures of drug use are the prevalence of use in the past 30 days of marijuana/hashish, cocaine, psychotherapeutic drugs (stimulants, sedatives, tranquilizers, analgesics), any drug, and any drug except marijuana. "Any drug" includes marijuana or hashish, inhalants, hallucinogens, cocaine, heroin, and psychotherapeutic drugs. The measures of cigarette smoking are the prevalence of any smoking (one or more cigarettes) and heavy smoking (smoking one or more packs of cigarettes per day) during the past 30 days.

Results

Alcohol Use

Table 2 presents standardized comparisons of the prevalence of any drinking and heavy drinking among military personnel and civilians by age group. Overall, military personnel are substantially more likely than civilians to drink and to drink heavily. This pattern of differences is consistent for males and females and for younger and older age groups, although in some cases differences have p values greater than 0.05. The prevalence of any drinking among all military personnel is about 8 percentage points higher than among all civilians, and the prevalence of heavy drinking among all military personnel is about 10 percentage points higher than that among all civilians.

Military/civilian differences in *any drinking* are more pronounced among women than among men. The rate for any drinking for women in the military is about 15 percentage points higher than the rate for civilian women; for men, the difference in military and civilian rates is about 7 percentage points. Younger military women are almost as likely as military men to drink and more likely than civilian men to do so.

Military men and women are about twice as likely as civilian men and women to be *heavy drinkers.* About one-third of younger military men drink heavily, a rate that is substantially higher than for any other group.

Drug Use

The prevalence of any drug use is substantially lower among military personnel than among civilians, as shown in Table 3. The rate of any drug use among all military personnel is about one-third that among all civilians, 8.4 percent versus 24.1 percent. Similarly, the rate of any drug use among military men is about one-third that among civilian men, whereas the rate of use among military women is about half that among civilian women. These tendencies for any drug use generally hold for specific drugs. Drug use among military women is about the same as among military men, whereas use among civilian women is substantially lower than among civilian men.

For most measures of drug use, the prevalence of any drug use is highest among civilian men, followed by civilian women, military men, and military women. The data in Table 3 also show

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**Wheeless SC, Shah BV: RATIO2 and RATIO3. Research Triangle Park, NC: Research Triangle Institute, 1982, unpublished paper.

TABLE 3—Standardized Comparisons of Drug Use in Past 30 Days Among Military Personnel and Civilians, by Age Group

	Age Group/Population					
	18–25 years		26–55 years		Total	
Group/Measure	Civilian	Military	Civilian	Military	Civilian	Military
Males	(N = 794)	(N = 1,641)	(N = 1,279)	(N = 5,877)	(N = 2,073)	(N = 7,518)
Marijuana	27.3	11.4***	16.9	1.8***	22.0	6.5***
Cocaine	9.7	4.6**	6.0	0.6***	7.8	2.6***
Psychotherapeutics[a]	8.9	4.8*	3.8	1.7*	6.3	3.2**
Any drug[b]	30.0	14.1***	20.2	3.3***	25.0	8.6***
Any drug except marijuana[c]	15.3	8.1**	10.1	2.2***	12.7	5.1***
Females	(N = 1,008)	(N = 210)	(N = 1,803)	(N = 343)	(N = 2,811)	(N = 553)
Marijuana	16.5	3.8***	6.3	1.3***	11.8	2.6***
Cocaine	6.1	3.3	1.8	0.3*	4.1	1.9
Psychotherapeutics[a]	3.6	6.1	3.3	2.7	3.5	4.5
Any drug[b]	19.9	10.3***	9.4	3.7**	15.0	7.2***
Any drug except marijuana[c]	9.2	8.3	5.1	2.7	7.3	5.7
Total	(N = 1,802)	(N = 1,851)	(N = 3,082)	(N = 6,220)	(N = 4,884)	(N = 8,071)
Marijuana	26.2	10.6***	16.0	1.8***	21.0	6.1***
Cocaine	9.3	4.5**	5.6	0.6***	7.5	2.5***
Psychotherapeutics[a]	8.3	4.9*	3.8	1.7*	6.0	3.3**
Any drug[b]	29.0	13.7***	19.3	3.3***	24.1	8.4***
Any drug except marijuana[c]	14.7	8.1**	9.7	2.3***	12.2	5.1***

Note: Estimates are percentages. Ns (in parentheses) show the number of cases on which weighted data are based. Civilian data have been standardized to the military data by age, education, race/ethnicity, and sex.

*Different from civilian at the .05 level.

**Different from civilian at the .01 level.

***Different from civilian at the .001 level.

[a] Stimulants, sedatives, tranquilizers, and analgesics.

[b] Nonmedical use one or more times of marijuana or hashish, inhalants, hallucinogens, cocaine, heroin, or psychotherapeutics.

[c] Nonmedical use one or more times of drugs listed in footnote b except for marijuana or hashish.

Civilian data source: 1985 National Household Survey on Drug Abuse.

Military data source: 1985 Worldwide Survey.

greater drug use among the younger age group than among the older age group (about 10 percentage points for any drug use). This finding holds for both military and civilian populations. Thus, those ages 18 to 25 are at considerably higher risk of drug use than those who are older.

Cigarette Use

Military personnel are in general substantially more likely than civilians to be smokers and heavy smokers, as shown in Table 4. For the total group, the prevalence of smoking is about 5 percentage points higher among military personnel than among civilians, and the prevalence of heavy smoking is about 8 percentage points higher. Differences among military and civilian women are more pronounced than among military and civilian men. The rates of any smoking and heavy smoking among military and civilian men differ by 4 and 7 percentage points, respectively, while the same rates among military and civilian women differ by 13 and 14 percentages points, respectively. Military women are especially likely to smoke and smoke heavily: they have the highest prevalence of smoking (45 percent) and the same rate of heavy smoking as military men (30 percent). Rates of heavy smoking are particularly high among older military men and women, over one-third.

Discussion

After controlling for key sociodemographic characteristics known to be associated with substance abuse, clear differences remain in rates of substance use between military personnel and civilians. Military personnel are, in general, less likely to use drugs than civilians and more likely to drink and drink heavily and to smoke and smoke heavily. A striking finding is the extremely high rate of heavy drinking among younger military men (about one-third).

The substance use patterns of men and women are also of note. Among civilians, women are less likely than men to use drugs, alcohol, or cigarettes. Among military personnel, women are also less likely to use alcohol than men; in contrast, military women are equally as likely as military men to smoke and use drugs.

The precise factors accounting for differences in substance use are unknown but may plausibly be associated with distinctive conditions in military life and in military policies and programs directed toward substance use. The military policy of zero tolerance for drug use, which began in 1981 and is supported by urinalysis testing, has been a significant deterrent. Applicants to military service are tested for drug use, and active-duty personnel identified as users are subject to discharge from the military. This policy has, no doubt, been associated with rapid declines in drug use among military personnel during the 1980s, declines that have been much more rapid than those among civilians.

Further support for the deterrent effect of urinalysis testing comes from a recent study of military applicants.[12] Data from urine tests assessing the prevalence of marijuana and cocaine use found that applicants who knew they would be tested for drug use were less likely to test positive than those who did not know. These results suggest that some drug users who know about the screening program make sure they are "clean" before the test and that others decide not to apply. The decision of some drug users not to apply for military service probably accounts for some of the difference in military and civilian drug use.

TABLE 4—Standardized Comparisons of Cigarette Use in Past 30 Days Among Military Personnel and Civilians, by Age Group

	Age Group/Population					
	18–25 years		26–55 years		Total	
Group/Measure	Civilian	Military	Civilian	Military	Civilian	Military
Males	(N = 794)	(N = 1,632)	(N = 1,279)	(N = 5,857	(N = 2,073)	(N = 7,489)
Any smoking	36.9	43.0*	43.2	44.7	40.1	43.9
Heavy smoking	17.1	25.3***	28.4	34.8**	22.9	30.2***
Females	(N = 1,008)	(N = 208)	(N = 1,803)	(N = 340)	(N = 2,811)	(N = 548)
Any smoking	33.4	46.9*	30.5	42.7*	32.1	45.0***
Heavy smoking	15.7	26.2*	16.4	34.4***	16.0	30.0***
Total	(N = 1,802)	(N = 1,840)	(N = 3,082)	(N = 6,197)	(N = 4,884)	(N = 8,037)
Any smoking	36.5	43.4*	42.1	44.5	39.4	44.0*
Heavy smoking	17.0	25.4***	27.3	34.8***	22.2	30.1***

Note: Estimates are percentages. Ns (in parentheses) show the number of cases on which weighted data are based. Civilian data have been standardized to the military data by age, education, race/ethnicity, and sex.
*Different from civilian at the .05 level.
**Different from civilian at the .01 level.
***Different from civilian at the .001 level.
Civilian data source: 1985 National Household Survey on Drug Abuse.
Military data source: 1985 Worldwide Survey.

The higher rates of drug use among civilians suggest that military policies and practices effectively deter drug use in the military. In contrast, the lower rates of alcohol use and cigarette use among civilians than among military personnel suggest that certain aspects of military life may foster use and/or that military policies and programs directed toward alcohol and cigarette use have not been effective. Concerted efforts to decrease alcohol and cigarette use have been more recent and perhaps less intense than the efforts against drug abuse. The findings suggest the need for the military to intensify efforts to eliminate alcohol abuse and eradicate cigarette smoking. Military programs directed toward alcohol abuse should be targeted primarily toward young men who have the highest rates of heavy use. Programs against smoking should target both men and women.

Recent military policies and programs take a more stringent stance against alcohol abuse and smoking. These include an integrated health promotion program, an antismoking campaign, and specific policies to deglamorize and to limit accessibility to alcohol. This intensification of efforts promises further reductions in substance abuse. Data from the 1988 Worldwide Survey** show progress since 1985, with continuing significant reductions in the unstandardized rates of the prevalence of drug use (8.9 percent to 4.8 percent), alcohol use (86.6 percent to 82.8 percent), and cigarette use (46.2 percent to 40.9 percent) among military personnel in the past 30 days. Standardized comparisons conducted with the 1988 National Household Survey data, which will soon be available for public use, will reveal whether these recent deceases in use by military personnel and related decreases among civilians result in a continuing divergence of rates of use between military personnel and civilians or whether the rates are becoming more similar. Comparisons of the two sets of standardized comparisons will suggest whether the intensified military efforts against alcohol abuse and smoking have been effective. ☐

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Risk Factors for Suicide Attempts Among Navajo Adolescents

ABSTRACT

Background: Rates of adolescent suicide in the United States are highest among Native Americans but little is known about risk factors for suicide attempts in this population.

Methods: To identify risk factors for self-reported suicide attempts by Navajo adolescents, we analyzed the 1988 Indian Health Service Adolescent Health Survey that was administered to 7,254 students in grades 6 through 12 on the Navajo reservation. The responses of students reporting a past suicide attempt were compared to others.

Results: Nearly 15 percent (N = 971) reported a previous suicide attempt; over half of those admitted to more than one attempt. Controlling for age, a logistic regression model revealed the following associations with suicide attempts: a history of mental health problems (OR = 3.2); alienation from family and community (OR = 3.2); having a friend who attempted suicide (OR = 2.8); weekly consumption of hard liquor (OR = 2.7); a family history of a suicide or attempt (OR = 2.3); poor self-perception of health (OR = 2.2); a history of physical abuse (OR = 1.9); female gender (OR = 1.7); and sexual abuse (OR = 1.5).

Conclusions: Efforts to prevent adolescent suicide attempts in this population should target individuals with those risk factors of the highest risk and prevalence of exposure. (*Am J Public Health*. 1991;81:870–874)

David C. Grossman, MD, MPH, B. Carol Milligan, RN, MS, CNM, and Richard A. Deyo, MD, MPH

Introduction

In 1987, over 2,100 youths under age 19 in the United States killed themselves, making suicide the third most common cause of death in the US adolescent population. Although suicide rates in the US population have been relatively stable, rates among youths have tripled over the past 30 years, from 4.5 per 100,000 in 1950 in 13.1 per 100,000 in 1986.[1] Among US ethnic groups, American Indians and Alaskan Natives have the highest rates.[2,3] The suicide rate among Native American adolescents in 1987 was 26.8 per 100,000, more than double the national rate for adolescents of all races.[4]

For every completed suicide, it is estimated that eight or more are attempted.[5] The 1987 National Adolescent Student Health Survey reported that 11.1 percent of males and 17.5 percent of females from a national random sample of eighth and tenth grade students had harmed themselves in a way that might have caused their death.[6] The prevalence of suicide attempts among Native Americans may even be higher. Among a cohort of American Indian high school students attending a boarding school in the southeast US, 23 percent had attempted suicide during their lifetime.[7]

This high rate of self-destructive behavior among Native American adolescents warrants further investigation and culturally appropriate intervention. The 1988 Navajo Adolescent Health Survey offered an opportunity to examine a set of traditional risk factors that are thought to be important in the general population. This population was chosen because it had the largest and most complete participation in the Indian Health Service Adolescent Health Survey conducted in 1988. The purpose of this study was to identify risk factors for self-reported suicide attempts among adolescent students in the Navajo Indian population.

Methods

Instrument

The source of data was the 1988 Navajo Adolescent Health Survey, a subset of the Indian Health Service Adolescent Health Survey. This 162-item survey was designed as a health promotion tool for use by the Indian Health Service and individual communities to identify and address problems in adolescent health. The original instrument was designed by Adolescent Health Program of the University of Minnesota[8] and adapted for use in Native American populations. After pilot testing, the survey was administered to Native American students throughout the United States. Content of the survey included personal data, information about the home and family, general health, nutrition, substance abuse, social concerns, experiences and attitudes regarding sex, mental health, and injury. The data from the Navajo Reservation represented more than 70 percent of the national sample.

Address reprint requests to David Grossman, MD, MPH, Department of Pediatrics, University of Washington, and Harborview Injury Prevention and Research Center, 325 Ninth Avenue, ZX-10, Seattle, WA 98104. Ms. Milligan is with the Maternal and Child Health Branch, Navajo Area Indian Health Service, Window Rock, AZ; Dr. Deyo is with the Departments of Medicine and Health Services, University of Washington and Seattle Veterans Affairs Medical Center. This paper, submitted to the Journal August 20, 1990, was revised and accepted for publication January 8, 1991.

TABLE 1—Crude and Adjusted Odds Ratios of Risk Factors for Suicide Attempts

Category/Risk Factor	Odds Ratios Crude	 Adjusted*	95% CI
Child Abuse			
Physical abuse	4.0	1.9	1.5–2.4
Sexual abuse	3.4	1.5	1.2–1.9
Demography			
Female gender	1.8	1.7	1.4–2.0
Community of residence	1.4 (Community "D")	NS	
Exposure to Suicide			
Family history of attempt/completion	3.9	2.3	1.6–3.2
Friend history of attempt	5.0	2.8	2.3–3.4
Friend history of completion	3.7	NS	
Family Structure			
Biological mother absent	.8	NS	
Biological father absent	1.3	NS	
Parents separated	1.5	NS	
One parent dead	1.4	NS	
Self-perception of Health Status			
Poor Health	4.4	2.2	1.3–3.8
Mental Health			
History of mental, behavior problems requiring professional help	3.7	3.2	2.2-4.5
Extreme alienation from family & community	5.0	3.2	2.1–4.4
Substance Abuse			
Weekly use of beer/wine	1.4	NS	
Weekly use of hard liquor	4.2	2.7	1.9–3.9
Mother's daily use of beer/wine	2.6	**	
Mother's daily use of hard liquor	4.3	**	
Father's daily use of beer/wine	2.6	**	
Father's daily use of hard liquor	3.6	**	
School Perception and Performance			
"I hate School"	2.3	NS	
Less Than "C" Grade Point Average	2.0	**	

NS: Not significant after adjustment
*Adjusted for age & all covariates
**"N" Too small to enter regression (see text)

Study Site/Subjects

The Navajo Reservation is the largest Indian nation outside of Alaska with an estimated population of 190,000. Schools on or adjacent to the reservation were encouraged to participate, although each school or individual student could decline. A volunteer sample of 47 schools representing grades six to twelve participated in the survey. All students in participating schools were asked to voluntarily respond to the anonymous survey during class time. There was no attempt to survey students absent on that day. The response rate exceeded 70 percent of enrolled students in virtually all participating schools to yield a usable sample of 7,241 responses, representing approximately one-third of all eligible students on the Navajo reservation.

Risk Factors

To determine risk factors for self-report of attempted suicide, we analyzed this survey using one question as the outcome, or dependent, variable: "Have you ever tried to kill yourself?" The independent variables consisted of 28 survey items in eight categories (Table 1) hypothesized to be underlying risk factors for suicide attempts. There were one to six survey items in each category. All of the items, except alienation, were measured by a single survey question.

Alienation was assessed using a nine item scale (Table 2) that explores the respondent's connectedness with the family and community. This scale had substantial internal consistency (Cronbach's alpha = 0.85) and construct validity when compared to other items of a similar nature in the survey.

Students reporting a suicide attempt were compared to those with no such report for each of the hypothesized risk factors. Chi-square statistics were used to establish significance in these comparisons and the significance level was adjusted to .0018 by the Bonferroni correction for multiple comparisons. Items that met this level of significance were entered in a multiple regression equation controlled for age and gender. Both forward stepwise and structured regression strategies were employed in this procedure. Using F tests, items that remained significant after adjustment for the other variables were entered into a logistic model. Estimates of adjusted odds ratios for each risk factor were derived from the coefficients of the logistic regression model. Categorical variables were divided into a series of dichotomous dummy variables; odds ratios presented for these variables represent comparison of the high versus low exposures. Data were analyzed with the SPSS-PC (SPSS Inc., Chicago, IL 60611) and EGRET (Statistics and Epidemiology Research Corp., Seattle, WA 98105) statistical packages.

Results

The sample was 51 percent female with a median age of 14.4 years and ninth

TABLE 2—Alienation Scale Items

How much do you feel:

1. Adults care about you?
2. School people care about you?
3. Your parents care about you?
4. Church leaders care about you?
5. Tribal leaders care about you?
6. Your family cares about your feelings?
7. People in your family understand you?
8. You and your family have lots of fun together?
9. Your family pays lots of attention to you?

5 point Likert Scale (Not at all—Very much)

1 2 3 4 5

Reliability (Cronbach's alpha) coefficient = 0.85

Mean score: 32.3 (SD 8.0)

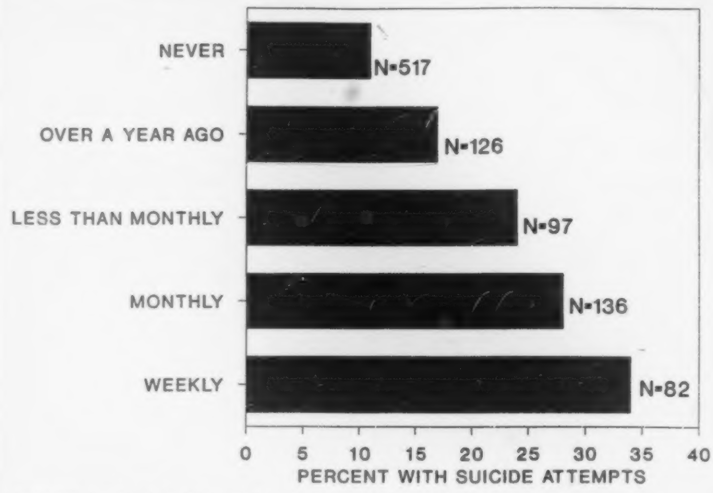

FIGURE 1—Suicide Attempts by Hard Liquor Use among Navajo Adolescents.

grade standing. Exposure to suicidal behavior among friends and family was common in this group of students. About 18 percent of the respondents replied that someone in their family had tried to kill or had killed themselves. Similarly, almost 18 percent had a friend who had attempted suicide and 9 percent had a friend who had completed suicide.

Of all respondents, 6,637 (92 percent) answered the question about suicide attempts; 971 (15 percent) admitted to a past attempt to kill themselves. Thirty-five percent of these attempts had occurred more than one year before the survey and 58 percent of the students had tried to kill themselves more than once. Thirty-six percent of suicide attempts were not reported to anyone and only 15 percent were brought to the attention of a medical professional.

In Table 1, the crude and adjusted odd ratios for hypothesized risk factors that were significant in distinguishing between respondents who had reported a previous suicide attempt, and those who had not, are shown.

In the final model, nine risk factors demonstrated a statistically significant and independent association with a history of suicide attempts. The two strongest associations were a history of a mental, behavioral, or emotional problem requiring professional help (OR = 3.2) and extreme alienation from family and community (OR = 3.2). Exposure to suicide completions and attempts by family and friends also shared strong associations with suicide attempts. Having a friend who previously attempted suicide (OR = 2.8) was more strongly associated with attempts by the respondent than having a family member either attempt or complete suicide (OR = 2.3). The consumption of hard liquor at weekly intervals was also significantly associated (OR = 2.7) with suicide attempts. The frequency of hard liquor consumption was linearly correlated with a lifetime prevalence of suicide attempts (Figure 1). Beer and wine consumption were not significant correlates at any frequency of use after adjustment for hard liquor use.

A self-perception of poor general health was also independently associated with a history of suicide attempts (OR = 2.2). The association of past physical abuse (OR = 1.9) with suicide attempts was marginally stronger than that of sexual abuse (OR = 1.5). Finally, the odds ratio for female gender remained significant after adjustment for covariates (OR = 1.7).

Several risk factors appeared to be associated with a past suicide attempt in the bivariate analysis but were not included in the model because of the many missing values. These variables included both the maternal and paternal consumption patterns of alcohol and the respondent's school performance as measured by the grade point average.

Risk factors which were significant in the bivariate analysis but did not remain significant after adjustment for other covariates, or after correction for multiple comparisons, included: community of residence, history of friend's suicide completion, all of the family structure variables, the weekly use of beer or wine, household density, religiosity of the respondent, the household presence of the natural mother, and an adoptive father or stepfather.

Discussion

To our knowledge, this is the first large controlled study of suicide attempts among Native American adolescents. The lifetime prevalence of self-reported attempts among Navajo students is high, but similar to that reported in the Minnesota and National Adolescent surveys.[6,8] The risk factors ascertained in this study of Native American youth are similar to those described for the general population.

An important finding of this study is the link between the suicide attempts of the respondents and those of their friends and family. Because prevalence of exposure was the highest for these risk factors and because of their strong association with suicide attempts, these risk factors have the greatest public health importance. Although bereavement following a family suicide is recognized to be a risk factor for suicidal behavior in older adults,[9] there have been few controlled studies documenting this risk in children and adolescents. The studies of adolescent suicide clusters, where a number of suicide deaths occur in a community over a short period, have been conflicting as to whether exposure to a friend's suicide is a risk factor.[10–12] But these studies have only evaluated deaths, not attempts, and have studied only suicide deaths associated with a suicide cluster.

Alienation from family and community was independently associated with a history of previous suicide attempt. This supports findings from numerous uncontrolled studies in other populations, including Durkheim's sociologic analysis of suicide.[13]

The association of alcohol and suicide attempts confirms previous findings and demonstrates new ones. Chronic and acute ethanol intoxication appear to play an important role in completed suicides of both White and Native Americans.[14,15] In contrast, there are few data on the role of alcohol in suicide attempts, nor on the differences between varying types of alcohol. After adjustment for beer and wine consumption, use of hard liquor remained the significant alcohol-associated variable related to suicide attempts. The linear relationship between frequency of hard liquor use and the lifetime prevalence of suicide attempts suggests that its consumption (or a close proxy) may play a causal role in suicide attempts.

The finding that past sexual or physical abuse is a risk factor for suicide attempts correlates with reports from other populations. Riggs, *et al*, recently reported similar findings for sexual and physical abuse in a smaller survey of Rhode Island students, although these variables were not adjusted for other behavior and life events.[16] Our study also indicated an interaction between the gender and abuse variables, demonstrating that sexual abuse of males and physical abuse of females are associated with higher risks than the converse situations. This has not been described in other studies.

Mental health problems were expected to be risk factors as their association to suicidal behavior has been well described.[17–19] Despite this, few controlled studies have confirmed this finding in the American Indian population. Manson's survey of Indian adolescent students from a southeastern boarding school demonstrated that those with a past suicide attempt had a higher level of depressive symptomatology.[7] A recent controlled study of adolescents presenting to an urban emergency room for suicide attempts indicated that they were more likely than controls to have a history of mental health care.[20]

There are several limitations to this study. Its cross-sectional design did not allow us to establish the temporal relationship between these risk factors and the suicide attempt. Mental health problems or alienation, may show associations because they may also be sequelae of an attempt. In addition, some associations may represent proxies for other variables not in the model. An example of a potential proxy might be child abuse, where the abuse history is important because it is symbolic of pervasive family dysfunction.

The survey did not measure the severity of the self-inflicted injury. Since only 15 percent of the attempts were reported to medical professionals, we infer that the unreported attempts represented gestures or minor injuries. Alternate interpretations are that potentially lethal attempts were only marginally successful, or that some of these unreported attempts were disguised as unintentional injuries, e.g. single occupant motor vehicle crash.

Because the study was conducted in school, it did not include those students absent on the day of the survey. Students who had been expelled, suspended, or had dropped out before the survey also were not represented. All of these groups may be at higher risk for many of these behaviors and life events. This may have led to an underestimate of the true lifetime prevalence of suicide attempts.

Although the use of anonymous self-reported questionnaires precluded corroboration of responses with existing data sources such as medical and social service records, its anonymity may have been an important safeguard of its validity. Since many of these behaviors are illicit or taboo, adolescents may be reluctant to disclose such information in a survey with identifiers. The validity of the self-reported information is supported by the similarity of these results to those from other studies.

Finally, this survey did not include some potentially important risk factors in this population. Exposure to traumatic deaths other than suicides and adaptability to a bi-cultural environment were not included but may also be important risk factors.[21,22]

The causal model of suicidal behavior is extremely complex and appears to vary between age and cultural groups. The causal model will be constructed using risk factors derived from controlled studies. Future preventive strategies should focus on identified risk factors that can be modified.[23] Furthermore, public health efforts to reduce suicide rates will be most successful if focused on those risk factors with the highest attributable risk. Future research efforts to determine the etiology of suicidal behavior in the American Indian and Alaskan Native populations should be directed toward carefully designed population-based case-control studies. ☐

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Job Stress Is "Epidemic," Costly to Business, Study Finds

One in three Americans seriously thought about quitting work in 1990 because of job stress, and one in three expects to "burn out" on the job in the near future.

A recent study also claims that workplace stress is causing far greater turnover, lower productivity, higher absenteeism, and more frequent health problems than was previously recognized. Employees who are overworked or have little control over their job are most likely to be burnout victims.

These are among findings from a study prepared by Northwestern National Life (NWNL) Insurance Company identifying the extent, causes, costs, and solutions related to workplace stress. Unlike many earlier studies, the NWNL research is believed to be the first to correlate causes and cures in all occupations and populations across the United States.

Any major change in the workplace affecting staffing is likely to speed up employee burnout, the research suggests. Significant burnout occurred in companies that had substantially cut employee benefits, changed ownership, required frequent overtime, or reduced the work force.

The findings are from interviews with a representative sample of 600 American workers. To help address the problem, NWNL also assembled a panel of stress and wellness experts to construct a workplace stress test that enables a company to quickly measure its work force stress and identify the consequences and solutions.

Among NWNL's key findings are the following:

- Seven in 10 American workers say that job stress is causing frequent health problems, and is making them less productive.
- Thirty-four percent of Americans thought seriously of quitting their job in 1990 because of workplace stress. Fourteen percent of all employees said workplace stress actually caused them to quit or change jobs in the last 2 years.
- Forty-six percent say their job is highly stressful. This is double the findings of a 1985 National Health Interview Survey that said 20% of Americans experience a high level of stress.
- Seventeen percent report higher absenteeism due to job stress.
- Fifty-three percent of Americans are required to work more than 40 hours a week very or somewhat often.
- Eighty-two percent of workers say victims of burnout deserve disability pay from their employer.
- Peggy Lawless, NWNL's research analyst, noted that job stress and burnout have a foothold in the US work force without regard to region, size of company, type of job, or age of the worker.

Lawless said the survey also shows that companies with supportive work and family policies, health coverage for mental illness and chemical dependency treatment, improved communication, and flexible work hours have almost half the burnout rate of companies that don't have such policies.

Lawless likened workplace stress in the 1990s to the lack of workplace safety in the 1960s. "Stress is less visible than coal dust or dangerous machines, but it's more prevalent and just as damaging to health and productivity."

The research results and the NWNL Workplace Stress Test are contained in a copyrighted report by NWNL called "Employee Burnout: America's Newest Epidemic." Copies can be obtained by calling 612/342-7137.

Patterns of Mental Health Utilization Among Island Puerto Rican Poor

Margarita Alegría, PhD, Rafaela Robles, PhD, Daniel H. Freeman, PhD, Mildred Vera, PhD, Aida L. Jiménez, MA, Carlos Ríos, MPH, and Ruth Ríos, MS

ABSTRACT

Background: This paper describes utilization of mental health services by poor Puerto Ricans living on the island. It examines the utilization rates, within health sectors, and settings for the provision of mental health services.

Methods: Data are based on an islandwide probability sample of 18- to 64-year-old respondents living in low socioeconomic areas. We assessed need with the Psychiatric Symptom and Dysfunction Scales.

Results: Approximately one-third of our study population (31.5%) met criteria for need. Of these, only 32% had received any mental health care in the past year. Need was significantly associated with use of physical or mental health services for mental health problems. We found those who needed services to be five times more likely than those who did not need services to have used one or both sectors of care at least once in the past year. Among the first group 21.8% used the physical health sector to deal with mental health problems in contrast with 17.9% who sought care in the mental health sector. In the physical health sector, subjects used the public and private settings equally. In the mental health sector, 70% of subjects used the public setting.

Conclusions: This suggests the nonpsychiatric physician as a main provider for mental health treatment. (*Am J Public Health*. 1991;81:875–879)

Introduction

We report the distribution of mental health services provided by the medical (physical) and mental health sectors to low income people living in Puerto Rico. The relation of sociodemographic characteristics with utilization of mental health services is examined. Utilization data obtained in this study provide information on the gap between the need for and provision of mental health care.

Researchers have studied the use of mental health and health services by various ethnic minorities.[1–7] Some studies have found underutilization of mental health services by Hispanics,[2–5] while others question this underutilization.[6–7] Discrepant findings have been attributed to treatment surveys versus community surveys, to the criteria used to define Hispanics, and to differences in the need for mental health services.

Previous studies[8–10] indicate that persons of lower socioeconomic status (SES) have a low rate of mental health service utilization. A recurring argument is that the poor hold attitudes and ideologies dissonant with scientific medicine.[11,12] Hoppe and Heller[13] suggest that lower-class populations distrust those outside of their circle of family and friends, having heavy involvement in familial and peer relationships. Differences in values and priorities placed on health across socioeconomic groups may better explain observed differences in the use of mental health and physical health sectors. For example, Ware and Young[14] reported that persons of lower SES value their physical health more than their mental health. Johnston and Ware[15] report that SES appears associated with somatization, recognition of illness symptoms, and behaviors in response to illness. Consistent evidence has shown that a substantial proportion of persons with emotional problems and mental disorders are treated in the physical health sector.[16–20] Hough, *et al*,[21] found that although Mexican Americans made fewer visits to mental health providers than their non-Hispanic counterparts, they did not differ in the proportion of visits to a physical health provider for mental health reasons.

This study concentrates on the population in the lowest socioeconomic strata since there is evidence to suggest this group is greatly exposed to economic strains, physical illness, and disabilities, known risk factors for mental health needs.[22,23] Little is known to explain the use of mental health services for individuals that belong to the lowest income strata.[24] In this paper, we examine the utilization of mental health services by this population on the island of Puerto Rico.

Methods

The target population of this study included the civilian, noninstitutionalized residents, ages 18 to 64, living in low so-

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From the Center for Evaluation and Sociomedical Research, University of Puerto Rico School of Public Health (all authors except Freeman); Dr. Freeman is with the Department of Community and Family Medicine, Dartmouth Medical School, Hanover, NH.

Address reprint requests to Margarita Alegría, PhD, Director, Center for Evaluation and Sociomedical Research, School of Public Health, Medical Sciences Campus, University of Puerto Rico, PO Box 365067, San Juan, PR 00936-5067. This paper, submitted to the Journal July 9, 1990, was revised and accepted for publication January 22, 1991.

cioeconomic areas of Puerto Rico. The population was sampled using a two-stage stratified cluster sample of the island. The strata were urban and rural, which economically defined sub-strata. The substrata or segments were classified as economically depressed based on an index of media house rent, family income, and house value. These segments were divided into clusters from which 178 were selected for our sample. The list of clusters defined a survey area which contained about 57 percent of the total island population living in poverty. The nonpoor living in these areas were surveyed but excluded from the analyses reported here. The poor people living in nonpoor areas were also excluded from the sample. The sample for this study was obtained using the sampling frame prepared by the Department of Labor and Human Resources, and based on the 1980 Census of Population and Housing as updated by new housing developments in 1986.

The probability sample of clusters yielded 2,682 housing units. Of these, 27.9 percent (748 housing units) were not eligible. Reasons for ineligibility included vacancies (292) and noneligible residents (456) 18 to 64 years of age.

One respondent in each household was randomly selected from among those eligible. The noninterview rate was 8.1 percent distributed as: 2.5 percent refusals; 5.4 percent persons unable to be contacted after three callbacks; and 0.3 percent other breakoff. A total of 1,777 respondents were successfully interviewed during February to October 1989, yielding a completion rate of 91.9 percent. We constructed a poverty indicator based on size of family and number of related children according to the criteria of the Department of Commerce, Bureau of the Census.[25] Of those respondents living in poor areas, 73.6 percent were assessed to be in poverty status according to the 1989 Census definition. Our analyses are limited to this subpopulation.

All analyses incorporate the underlying survey design. Estimation and testing was performed using the program PC Carp.[26] Group differences in sociodemographic characteristics for utilizers of the physical health sector and the mental health sector for mental health services were tested using Z scores without multiple comparison adjustments. Sociodemographic differences with regard to service sectors used were tested with analyses of variance. Differences are not discussed unless the p value was less than 0.05. In all comparisons we will discuss the weighted sample so our inferences will apply to the Puerto Rican poor population residing in low socioeconomic status areas.

Measures

Respondents were interviewed by lay interviewers who gathered data on sociodemographics, psychiatric symptomatology, physical health, and utilization of physical health and mental health services. Questions on utilization of physical and mental health services for treatment of mental health problems were based on those of the National Institute of Mental Health (NIMH) Epidemiologic Catchment Area studies.[20] For the purposes of this paper, health care use will be limited to the formal system of care, recognizing that the informal system of care (e.g. clergy, friends, family) may also be an important link for explaining how Puerto Ricans cope with their emotional problems.

A mental health care visit is defined as a contact with any type of formal health care provider for a mental health problem. Included in the definition are visits to any type of formal health facility, such as a community health center, or to a physical health provider to deal with mental health problems. The mental health related contacts are examined in the physical health sector and the mental health sector. The physical health sector represents visits to a non-psychiatric physician for mental health problems. The mental health sector corresponds to visits to psychiatrists, psychologists, or any mental health treatment program regardless of whether there was a mental health reason for the visit.

Psychiatric symptomatology was assessed by the Psychiatric Symptom and Dysfunction Scales (PSDS). It consists of five subscales: anxiety, depression, psychosocial dysfunction, cognitive impairment, and general psychopathology. The items for these subscales were selected from the psychiatric, epidemiologic, and clinical research literature for the Florida Health Study.[27-30] Psychiatric symptoms and nondiagnosed psychiatric disorder were the criteria employed to determine need for mental health services. The scale provides a normative distribution of psychiatric symptomatology and dysfunctions in the population. They do not diagnose individuals. The PSDS has been found to be a valuable and adequate measure to identify differential levels of mental health need within the general and patient populations, including Hispanics and other ethnic groups on the mainland.[28,29] Persons scoring with moderate or high symptoms on 3 to 5 scales are included in the need category[2] for mental health services and are termed "needers."

These scales were adjusted for number of chronic illnesses, since some PSDS scores increased with the number of chronic illnesses. Specifically, psychosocial dysfunction, anxiety, and depression scores were reduced two, two, and three points, respectively, for every chronic health illness reported by the respondent. Cognitive impairment and general psychopathology scores were reduced by one point for each chronic physical illness. The actual construction of the scale permits no probing to differentiate whether the apparent psychiatric symptom (e.g. sweaty hands) has some physiological origin. Therefore, adjusting the scales in this respect reduces the upward bias due to poor physical health.

Results

The sociodemographic characteristics of respondents who utilized the mental or physical health sectors for mental health reasons are presented in Table 1. In general, utilization of physical health services did not vary with age. Use of mental health services was significantly higher for those 45–54 years old ($p < .05$) compared to those 18–24 and 35–44 years of age. A higher percentage of respondents with less than 10 years of school education than those with higher scholastic achievement reported use of physical health services for mental health reasons during the last year. Use of mental health services was significantly lower among respondents with some college education ($p < .01$) in comparison to those with a lower educational level. The retired or incapacitated were more likely than others to use the mental health sector or both sectors to deal with emotional problems. Sex, marital status, insurance, and zone of residence did not show a statistically significant association with health services sectors used for mental health reasons.

The utilization of the health and mental health care sector for mental health reasons during the previous year is shown in Table 2. As expected, needers are five times as likely to use one or both sectors of care at least once in the past twelve months for mental health problems (31.5 vs 5.8 percent). This discrepancy is even greater when we consider the use of the general health sector at least on one occasion (21.8 vs 3.7 percent) during the last year. These visits were identified by the respondents as health encounters to deal with mental health problems. The utilization of the mental health sector showed a higher prevalence of mental health visits by needers than by non-needers (17.9 vs 2.3 percent). Among users, the intensity varied by sector and need status. Needers averaged 7.3 (±0.9) visits to the physical health providers while non-needers averaged only 3.5 (±0.6) visits ($p < 0.001$). In contrast, there were no significant differences in the mental health sector utilization visits by needers (9.6, ± 1) versus 11.1 ± 2.7 visits by non-needers ($p > 0.25$).

Nearly 13.6 percent of needers of mental health services are receiving care exclusively in the physical health sector (row 4, Table 2). Furthermore, there is a small group of needers who use both sectors for mental health reasons (8.2 percent). This dual usage of mental health services in both the physical and mental health sector rarely occurs among non-needers (0.3 percent).

Table 3 presents the utilization during the last year of the public and private settings for a mental health problem by need status. Utilizers of exclusively the physical health sector were equally likely to choose a private or public setting for mental health care. In contrast, utilizers of exclusively the mental health sector employed by a 3 to 1 margin the public over the private setting. Approximately one-half of the users of public mental health services reported using mental health centers (data not shown).

TABLE 1—Sociodemographic Characteristics of the Sample Below Poverty Level by Health Sector Utilized for Mental Health Reasons During the Last Year

	Utilizers of Services for Mental Health Problems			
Sectors:	Physical Health	Mental Health	Both	Non-users
Population N	46,980	31,590	19,710	609,390
Sample n	96	59	42	1,111
	% (S.E.)	% (S.E.)	% (S.E.)	% (S.E.)
Age (years)				
18–24	16.7 (5.0)	12.0 (3.9)	5.5 (3.3)	25.9 (1.5)
25–34	17.2 (3.6)	21.4 (5.2)	12.3 (4.4)	26.2 (1.5)
35–44	21.3 (4.7)	16.2 (5.0)	24.7 (7.7)	21.0 (1.4)
45–54	25.9 (5.3)	32.5 (7.4)	28.8 (8.2)	13.2 (1.2)
55–64	19.0 (4.0)	17.9 (5.7)	28.8 (6.9)	13.6 (1.2)
Education (years)				
0–6	33.9 (5.6)	35.9 (7.6)	41.1 (9.3)	24.7 (1.4)
7–9	26.4 (6.0)	15.4 (5.0)	23.3 (7.1)	20.4 (1.2)
10–11	10.9 (3.3)	18.8 (5.8)	2.7 (1.9)	9.9 (1.0)
12	16.7 (4.3)	25.6 (6.4)	17.8 (6.2)	28.0 (1.4)
13+	12.1 (4.3)	4.3 (2.4)	15.1 (5.8)	17.0 (1.3)
Sex				
Male	33.9 (5.5)	41.9 (1.1)	57.5 (1.2)	39.8 (1.5)
Female	66.1 (5.5)	58.1 (7.6)	42.5 (8.4)	60.2 (1.5)
Marital Status				
Not Married	43.1 (5.8)	46.2 (6.8)	37.0 (7.6)	40.4 (1.8)
Married	56.9 (5.8)	53.8 (9.6)	63.0 (1.1)	59.6 (1.8)
Employment				
Employed	30.1 (6.1)	11.5 (4.3)	12.5 (7.1)	40.7 (2.0)
Unemployed	39.1 (7.0)	22.1 (5.9)	9.4 (3.7)	34.1 (1.9)
Retired or Incapacitated	21.1 (4.9)	56.7 (8.1)	71.9 (8.2)	10.0 (1.1)
Student	9.8 (5.7)	9.6 (3.9)	6.3 (4.0)	15.3 (1.5)
Insurance				
No	85.6 (3.9)	84.6 (7.3)	69.9 (1.1)	87.6 (1.3)
Yes	14.4 (3.9)	15.4 (5.2)	30.1 (7.9)	12.4 (1.3)
Zone of Residence				
Urban	46.6 (5.6)	57.3 (7.0)	53.4 (9.1)	55.7 (1.9)
Rural	53.4 (5.6)	42.7 (10.0)	46.6 (1.3)	44.3 (1.9)

TABLE 2—Utilization in the Past Year of Physical Health and Mental Health Services for Mental Health Reasons

		Total Sample	Needers	Non-needers
Respondents with one or more visits to:	Population N	707,670	222,750	484,920
	Sample n	1,308	413	895
		% (S.E.)	% (S.E.)	% (S.E.)
Any Sector		13.9 (1.1)	31.5 (2.6)	5.8 (0.8)
Physical Sector⁺		9.4 (0.9)	21.8 (2.4)	3.7 (0.6)
Mental Health Sector⁺		7.3 (0.8)	17.9 (1.9)	2.3 (0.5)
Only the Physical Health Sector		6.6 (0.7)	13.6 (2.0)	3.5 (0.6)
Both Physical Health and Mental Health Sector		2.8 (0.5)	8.2 (1.5)	0.3 (0.2)

Numbers in parentheses indicate standard errors.
⁺Does not exclude the use of other sectors.

Discussion

In contrast to other studies,[18,31,32] poor women in our sample were as likely to receive mental health care as men; elevated rates of unemployment (38 percent) among the poor may be a contributing factor to this finding. Unemployment also may be related to diminished social standing and vulnerability to psychological distress in men, leading them to seek mental health services.

Insurance did not appear to facilitate entrance into mental health care. This could be an artifact of Medicaid eligibility, since it was reported by our respondents as noninsurance. Medicaid operates as a centrally administered program by the Department of Health, whereby participants have free access to public health facilities. Public health and mental health services have an open door policy which may render those services accessible to the "uninsured."

Total mental health visits among poor Puerto Ricans (31.5 percent in our study sample) were twice as high as those reported in the ECA sites.[20] Economic factors and impoverished social conditions may contribute to a higher need for mental health services.

The finding that almost 22 percent of the needers went to professionals who are not specialists in mental health suggests a possible link between mental health disorders and physical illness, or difficulties in discriminating psychiatric symptoms from somatic complaints as a cultural expression of discomfort. Findings by

TABLE 3—Utilization of the Physical Health and Mental Health Sectors by Setting and Need Status (previous year) for the Sample Below Poverty Level

Sector		Total	Needers	Non-needers
	Population N	98,280	70,200	28,080
	Sample n	197	137	60
All Sectors		% (S.E.)	% (S.E.)	% (S.E.)
Private setting*		51.1 (3.9)	52.3 (4.7)	48.1 (7.5)
Public setting		48.9 (3.9)	47.7 (4.7)	51.9 (7.5)
	Population N	46,980	30,240	16,740
	Sample n	96	59	37
Physical Health All Sectors		% (S.E.)	% (S.E.)	% (S.E.)
Private*		55.7 (5.8)	53.6 (7.5)	59.7 (8.9)
Public		44.3 (5.8)	46.4 (7.5)	40.3 (8.9)
	Population N	31,590	21,600	9,990
	Sample n	59	40	19
Mental Health		% (S.E.)	% (S.E.)	% (S.E.)
Private setting*		30.8 (6.5)	32.5 (8.0)	27.0 (10.5)
Public setting		69.2 (6.5)	67.5 (8.0)	73.0 (10.5)
	Population N	19,710	18,360	1,350
	Sample n	42	38	4
Both Physical and Mental Health		% (S.E.)	% (S.E.)	% (S.E.)
Private*		72.6 (7.8)	73.5 (8.0)	60.0 (24.0)
Public		27.4 (7.8)	26.5 (8.0)	40.0 (24.0)

Note: Numbers in parentheses indicate standard errors.
*Private includes both private and combined settings.

Wells, *et al*,[22] suggest that mental health visits to general health providers commonly involve the treatment of both physical and mental health symptoms. An alternative explanation of this phenomenon is that contacts with the general health sector fulfill additional treatment needs, including social support and advice.

Patients' ambiguous definitions of psychiatric symptoms mixed with physical ailments may deter early diagnoses of mental health symptoms by the nonpsychiatric doctor or lead to misdiagnosis. Our study sample may reflect a tendency toward somatization. Economic strains combined with frequent unemployment as well as constant exposure to crime and violence make more flexible community standards of what constitute mental health problems. Such conditions could legitimize feelings of anxiety, depression, or sadness. The social acceptance of psychiatric symptoms may deter family, friends, and health professionals from referring needers to specialized mental health services.

This may be a consequence of the nonpsychiatric physician being gate-keepers of mental health care. Less than one-sixth (15 out of 103) of the specialty contacts were actually referred by physical health providers. Lack of referrals by nonpsychiatric physicians may be a major barrier to access to mental health services. Limited training in recognizing mental health symptoms by nonpsychiatric physicians or a negative attitude toward mental health care may impede referral.

Limitations of this study should be noted. The population surveyed cannot be assumed to be representative of all poor respondents in Puerto Rico or poor Puerto Ricans in the mainland. Poor persons not residing in low income areas of the island, as well as those in institutions (such as prison) or living in the streets, were excluded from this study. These findings cannot be extrapolated to poor Puerto Ricans in the mainland. Regional differences in Medicaid benefit coverage or local availability of services may influence the utilization rates for mental health services. Language barriers may also prevent access to services outside the island.

We cannot confirm self-reported utilization information provided by our respondents, therefore the validity of the information is not measured. It is probable that our measure of need (PSDS) identifies only certain areas of need for mental health services by the poor. We found that the PSDS scales do not detect as needers a high percentage of individuals who fulfill DIS criteria for alcohol abuse and dependence. However, we found a high concordance in the identification of needers by the CES-D, self-perception of mental health, and the PSDS.

Finally, we made empirical adjustments for the upward bias of poor physical health on the PSDS scale scores, although such empirical adjustment could underestimate actual need for mental health services among the poor. In our estimates of need, we sought to identify need for mental health services not due to physical health problems. The reliability of utilization data could be affected by elapsed time between actual use of services and the interview. Data from Kelly, *et al*,[33] suggest that reliability of medical information was not materially affected by the time interval between interviews. However, it was generally better when the time elapsed was less than one year, as was the case in our study.

In summary, the high utilization rate identified in our study group might be related to the organization of the public health care delivery system. The system in Puerto Rico is a vehicle of access to care for all in need, thus the coverage of the population is not constrained by the Medicaid income threshold standards as in the mainland United States. This broader coverage could be instrumental in explaining utilization patterns of Puerto Rico's low income population when compared to their counterparts on the mainland. Such comparisons can help assess the effects of organizational factors that influence or inhibit contact between low socioeconomic groups and mental health professionals. ☐

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A Prospective Study of Hospitalization with Gallstone Disease Among Women: Role of Dietary Factors, Fasting Period, and Dieting

Rosely Sichieri, MD, PhD, James E. Everhart, MD, MPH, and Harold Roth, MD

ABSTRACT

Background: Dietary risk factors for the development of gallstones have not been clearly established. We analyzed data from a population-based prospective study to determine dietary risk factors for hospitalization with gallstone disease.

Methods: We evaluated the role of dietary constituents, fasting, and dieting on subsequent hospitalization with gallstone disease among 4,730 women, ages 25 to 74 years, who participated in the first follow-up of the first National Health and Nutrition Examination Survey. Baseline dietary variables were established through a 24-hour dietary recall and a medical history. Proportional hazards models were used to calculate the effects of dietary variables while controlling for baseline risk factors.

Results: After an average of 10 years follow-up, gallstone disease was confirmed by hospital records among 216 women who denied gallstone disease at the baseline examination. The hazard rate of hospitalization with gallstone disease increased with increasing overnight fasting period and with dieting. Intake of fiber showed a small protective effect. The effect of energy intake was significant only among women younger than age 50 years at baseline. Results were not affected by adjustment for known risk factors for gallstone disease or other dietary factors.

Conclusion: A long overnight fasting period, dieting, and low fiber intake may increase the risk of hospitalization with gallstone disease. (*Am J Public Health*. 1991;81:880–884)

Introduction

Cholecystectomy is one of the most common surgical procedures in the United States, ranking fifth among all nonobstetric, therapeutic hospital procedures. In 1987, cholelithiasis, the leading reason for cholecystectomy, was the first listed diagnosis for 522,000 hospital discharges.[1]

Differences in diet are an appealing explanation for the wide variation in the occurrence of gallstones in the United States and internationally.[2,3] Numerous clinical and epidemiological studies have examined potential dietary risk factors for gallstones.[4–14] Such risk factors have included, among others, cholesterol,[4–6] linoleic acid,[5–7] oleic acid,[5–7] total carbohydrates,[6,8] fat,[6,8] fiber,[8–10] calcium,[11] and energy intake.[12,13] However, the results of these studies have not been consistent. Most of the epidemiological studies have used cross-sectional or case-control designs. One prospective study of energy intake found a positive association[12] with gallstones and another found no relation.[13] After examining more than 50 foods and nutrients the only association with diet in a recent prospective study of 88,373 middle age women was a negative one with vegetable fat and protein.[14] Clinical studies have suggested that certain dietary patterns, such as weight loss dieting and prolonged fasting may increase gallstone risk.[15–17]

The first National Health and Nutrition Examination Survey (NHANES I) collected dietary information. Follow-up interviews averaging 10 years later allowed prospective evaluation of dietary factors on subsequent disease. We tested specific hypotheses regarding the roles of dietary constituents, fasting, and dieting on the occurrence of hospitalization with gallstone disease in the longitudinal study of this population.

Methods

Subjects and Study Design

Baseline data, including dietary information, were collected in 1971–75 among a national probability sample of the United States population.[18,19] Follow-up of this cohort occurred in 1982–84 for those ages 25 to 74 years at the baseline examination.[20] The current study is limited to women because of the small number of cases of gallstone disease among men. Analysis was restricted to 4,730 Black and White women after excluding loss to follow-up (1,032 women), self-reported gallstones at baseline (591 women), unavailability of confirmatory hospital records for self-reported diagnosis of gallbladder disease during the follow-up (363 women), and lack of baseline information on gallstone disease (21 women), diet (65 women) or body mass index (two women).

Hospitalization with gallstone disease in the follow-up was coded according to the Ninth Revision of the International Classification of Diseases with Clinical Modification (ICD-9-CM). Cases had hos-

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From the National Institutes of Health, National Institute of Diabetes and Digestive and Kidney Diseases, Epidemiology and Data Systems Program. Address reprint requests to James E. Everhart, MD, MPH, NIH/NIDDK, Federal Building, Rm 106, 7550 Wisconsin Avenue, Bethesda, MD 20892. Dr. Sichieri is also affiliated with the Centro de Ciências Biológicas e de Saúde Universidade Estadual de Maringá, Paraná, Brasil. This paper, submitted to the Journal June 20, 1990, was revised and accepted for publication January 23, 1991.

pital discharge diagnoses of cholelithiasis (ICD-9-CM 574) (193 cases) or cholecystitis (ICD-9-CM 575.0–575.1) plus cholecystectomy (ICD-9-CM, procedure 51.22) (23 cases). The first date of hospitalization was regarded as the date of occurrence of clinical gallstone disease. Remaining for analysis were 216 women with hospital confirmed gallstone disease and 4,514 women without evidence of gallstones. A more detailed description of the study cohort and diagnostic criteria has been published.[21]

Variables and Measurements

All independent variables were measured at the baseline examination. Nutrient and energy intake were estimated from a 24-hour recall interview.[18] Participants were asked what they ate and the respective time, by hour and minute, for the 24-hour period beginning at midnight of the day before the interview. Based on these times two variables were created: number of meals and snacks per day, and the duration of overnight fasting period. Total number of intakes at least one hour apart equaled the number of meals and snacks. Overnight fasting period was estimated by the interval between the time of the last meal or snack and the first meal or snack ((24 hours − time of last meal) + time of first meal), using a 24-hour clock. Persons who ate only once during the 24-hour period were assigned a fasting period of 24 hours. Daily nutrient and energy intake were determined using standardized food composition tables applied to the foods reported in the dietary interview.[22] Fiber consumption was based on a provisional table of dietary fiber composition of foods.[23]

Self-reported alcohol consumption, recreational and nonrecreational physical activity, and dieting were ascertained from the baseline medical history interview. Dieting was based on the question "Are you on a special diet?"

Data Analysis

Dietary variables analyzed were those shown in Table 1 plus dieting. All of these variables have been postulated as having potential gallstone promoting activity.[4–13] Overnight fasting period was analyzed both as a continuous variable and as indicator variables (less than eight hour, eight to less than 10 hours, 10 to less than 12 hours, 12 to less than 14 hours, and at least 14 hours) to determine if there was a particularly increased risk with long fasting as has been suggested from clinical studies.[24]

TABLE 1—Medians, 25th and 75th Percentiles for Dietary Risk Factors for Women at the Baseline Examination, NHANES I, 1971–75

Factors	25th	Median	75th
Overnight fasting period (hr)	10.0	11.5	13.5
Number of meals (per day)	3.0	4.0	5.0
Energy (kcal)	1062	1412	1814
Carbohydrates (g)	115	160	210
Fat (g)	39	56	77
Saturated fat (g)	13	19	28
Oleic acid (g)	14	21	30
Linoleic acid (g)	3.0	5.4	8.9
Cholesterol (mg)	135	243	423
Calcium (mg)	287	503	790
Fiber (g)	4.8	7.5	11.2

In accordance with a previous analysis, non-dietary variables included were age in years and body mass index (weight in kg/height in m^2) as continuous variables, Black or White race, and parity categorized as at least four live births or less than four live births.[21] We also included as covariates alcohol consumption (g/day)[25] as a continuous variable and an index of physical activity. This last variable was constructed by summing the self-reported recreational and nonrecreational activities, each categorized in three levels.

Multivariate analysis used proportional hazards survival models with confirmation of the proportional hazards assumption of relatively constant rate ratios during the follow-up.[26] The effect of each dietary variable on the risk of gallstone disease was evaluated controlling for the non-dietary covariates: age at the baseline examination, body mass index (BMI), parity, race, quadratic terms for age and for BMI (age times age and BMI times BMI), and an interaction term for age and race (age times race). Interaction terms between the dietary variables and age, race, and BMI were examined for each dietary factor and were kept in the model if they improved significantly (p value < 0.05) the likelihood ratio of the model.[27] If an age interaction was present, the analysis was divided between persons less than age 50 and at least 50 years of age to simplify the models. Adjustment for energy consumption was made by including the 24-hour energy intake. For models including all ages, an interaction term of energy intake and age was added.

There were 65 women with no information on dieting and very few women (fewer than 13) with missing values for the other variables. Univariate analyses included all available information and multivariate analysis excluded persons with any missing values.

Results

Eighty-two percent of the women reported that what they ate during the 24-hour recall was the way they usually ate. Medians and the 25th and 75th percentiles are shown for each dietary variable at baseline (Table 1). Fifteen percent of women reported being on a special diet.

Survival analysis showed that overnight fasting, dieting, and fiber consumption were significantly related to gallstone disease (p value ≤ 0.03) when each dietary factor was adjusted for body mass index, body mass index^2, age, age^2, race, race times age, and parity. Number of meals, and total carbohydrate intake were marginally related to gallstone disease ($0.05 <$ p value < 0.10). Age influenced the effects of energy and fat intake (Table 2). Hazard rate ratios and 95 percent confidence intervals unadjusted and adjusted for energy intake are shown in Table 3 for those dietary factors with p values ≤ 0.10 in Table 2. Since the effects of energy intake and fat consumption were dependent on age, these variables were evaluated in separate models for women under and over age 50 years at the baseline examination. Energy adjustment reduced the hazard rate ratio for all variables.

When overnight fasting period was divided into five discrete periods the trend for an increased hazard ratio with increased overnight fasting was similar to the exponential model with overnight fasting as a continuous variable. Risk was greatest for 14 or more hours fasting, intermediate for 8–14 hours, and the least for fasting less than eight hours (Figure 1).

Additional analyses that included the significant dietary predictors of gallstone disease (dieting, overnight fasting period, fiber and energy intake), alcohol intake, and physical activity plus the other non-

TABLE 2—Hazard Coefficients[*] and Standard Errors (SE) for Dietary Factors for Hospitalization with Gallstone Disease for Women, NHANES I Epidemiologic Follow-up Study, 1971–84

Factors	Hazard Coefficient	SE	p value
Overnight fast period (hr)	0.078	0.024	0.003
Number of meals (per day)	−0.117	0.065	0.07
Dieting (yes/no)	0.513	0.158	0.001
Energy (kcal)	−0.0012	0.00041	0.02
Energy × age	0.00002	0.000007	0.007
Carbohydrates (g)	−0.0017	0.0010	0.08
Fat (g)	−0.0223	0.0080	0.006
Fat × age	0.0004	0.0001	0.01
Saturated fat (g)	−0.0057	0.0058	0.32
Oleic acid (g)	−0.0070	0.0057	0.22
Linoleic acid (g)	−0.0244	0.014	0.10
Cholesterol (mg)	−0.00016	0.00029	0.58
Calcium (mg)	−0.000006	0.0001	0.97
Fiber (g)	−0.029	0.014	0.03

[*]Each dietary variable evaluated individually adjusting for body mass index, body mass index^2, age, age^2, race, age times race, and parity.

TABLE 3—Hazard Ratios (HR)[*] and 95% Confidence Intervals (CI) for Dietary Factors for Hospitalization with Gallstone Disease among Women, NHANES I Epidemiologic Follow-up Study, 1971–84

		Unadjusted for energy intake		Adjusted for energy intake	
	Age	HR	95% CI	HR	95% CI
Energy (kcal)[†]	<50	0.69	0.53, 0.88		
	≥50	1.08	0.85, 1.36		
Overnight fasting period (hr)	all	1.33	1.10, 1.60	1.29	1.09,1.53
Dieting (Yes/No)	all	1.67	1.21, 2.23	1.66	1.21, 2.26
Number of meals/day	all	0.79	0.61, 1.02	0.83	0.64, 1.08
Carbohydrates (g)	all	0.85	0.70, 1.02	0.92	0.74, 1.34
Fat (g)[†]	<50	0.73	0.56, 0.95	1.08	0.62, 1.86
	≥50	1.07	0.88, 1.29	1.04	0.67, 1.60
Linoleic acid (g)	all	0.86	0.73, 1.02	1.10	0.90, 1.33
Fiber (g)	all	0.82	0.69, 0.97	0.85	0.70, 1.02

[*]All variables also adjusted for age, age squared, race, body mass index, body mass index squared, parity, and age × race. Models including all ages were also adjusted for energy × age. Except for diet, which was dichotomous, hazard ratios were obtained from models using continuous variables. Calculations were based on:

$$HR = e^{\text{hazard coefficient} \times (\text{75th–25th percentiles})}$$

[†]Variables with age interaction were analyzed by separate models for age groups: < 50 and ≥ 50 years.

dietary covariates specified in Table 3 did not substantially change any of the results (data not shown). Also, a model using as cases all women who reported hospitalization for gallbladder disease whether or not records were available gave essentially the same results (data not shown).

Discussion

In this prospective study of hospitalization with gallstone disease, a long overnight fasting period and dieting emerged as the two most important dietary risk factors for hospitalization with gallstone disease. Few studies have examined these risk factors. In two small, cross-sectional studies a positive association was found between duration of overnight fasting and gallstones.[11,28] A Danish population-based study of ultrasound diagnosed gallbladder disease found a history of "slimming treatments" associated with gallstones in men but not in women.[29] It was not known in the current study what proportion of women reporting a special diet were losing weight.

Long fasting period and dieting may act through diminished gallbladder motility, changes in the bile composition, or both to increase the risk of gallstones. Diverse conditions such as pregnancy,[30] parenteral nutrition,[31] diabetes,[32] and Chagas Disease[33] decrease gallbladder emptying and are considered risk factors for cholelithiasis. A clinical study of eight patients found that gallbladder contractions were reduced at mealtime during weight reduction, and the effect was more marked during prolonged fasting.[34] In addition, saturation of bile with cholesterol increases with fasting[24,35] and with weight loss.[36]

A recent study suggests that increased meal frequency may have a beneficial effect on health.[37] We did not find a strong relationship of number of meals and hospitalization with gallstone disease, but did find that a prolonged overnight fasting period increased the risk. Although fasting period was not a variable conceived when the original data were collected and edited, it is plausible that the times for meals and snacks reported on the 24-hour recall may represent usual eating times because of the regularity of daily activities for most persons. Longest fasting period based on 24-hour recall was also used in a Swedish study of diabetes,[38] but, as far as we know, it has not been validated, although 82 percent of the NHANES I study cohort considered the 24-hour recall their usual way of eating. Also, the increased risk after more than 14 hours of fasting resembles findings of bile composition studies in which lithogenicity of bile occurred only after a long period of fasting. Williams, *et al*, studied women without gallstones and found that saturation of bile with cholesterol was present in 4.5 percent after nine hours of fasting and in 54.5 percent after 16 hours of fasting.[24]

For the analyses based on the dietary composition, associations of gallstone disease hospitalization were found with low energy consumption among the younger group of women and women with low fiber intake. Other epidemiological studies have found negative,[8] positive,[12] and no relationship[13] between energy intake and gallbladder disease. The study that found a positive relationship was most similar to the current study, being a large, prospective study with extensive dietary data and an almost similar case definition. The negative association for energy intake that we found among younger women could be the result of a bias in the dietary recall or of differential energy expenditure between cases and noncases. Another possibility is

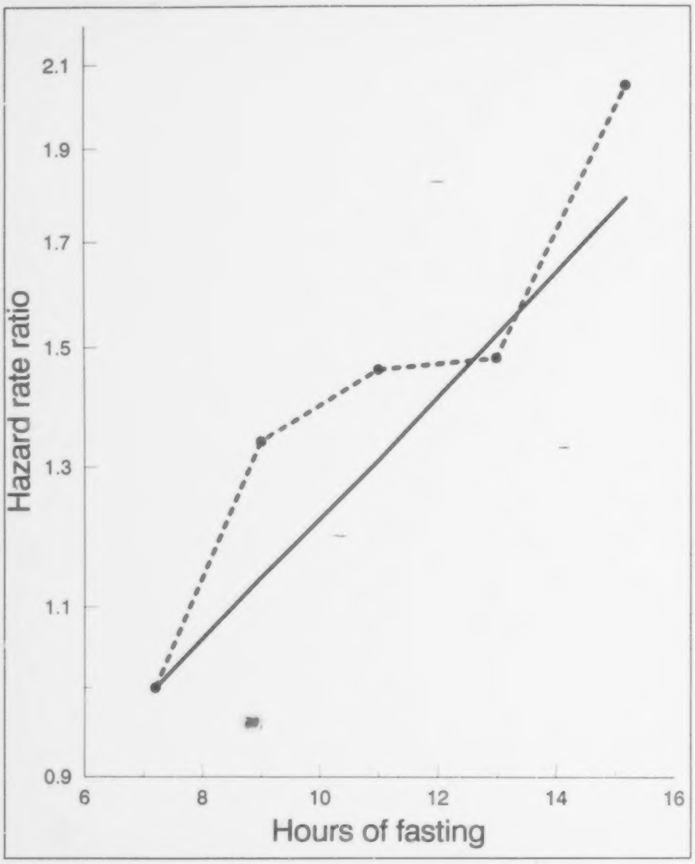

FIGURE 1—Duration of Overnight Fasting Period and Subsequent Hospitalization with Gallstone Disease in Women, NHANES I Epidemiologic Follow-up Study, 1971–84

The solid line represents the hazard rate ratio for overnight fasting period modeled as a continuous variable, compared with fewer than eight hours. Each point on the dashed line is the hazard rate ratio for ranges of overnight fasting period (8–10, 10–12, 12–14, and at least 14 hours) compared with fewer than 8 hours.

that women who have a propensity to form gallstones could have a higher metabolic efficiency and therefore lower energy intake that is independent of BMI.

There are also conflicting reports on the role of dietary fiber in gallstone disease.[9,10] Fiber is a convenient term for heterogeneous groups of poorly digested plant components whose nutritive functions are not well characterized. Colonic metabolism of fiber may affect bile secretion[39] and bind with bile acids in the gastrointestinal tract.[40]

Our results could not be explained by selective exclusion of cases without hospital records because analyses of all reported hospitalization for gallbladder disease, independent of availability of hospital records, showed similar results. Also, the prospective data collection eliminated the possible recall bias that is an unavoidable problem in case-control studies. Biased ascertainment could have occurred in this study if the risk factors analyzed increased the risk of gallstone symptoms rather than increase the risk of gallstones. However, a population-based study using ultrasonography showed that the frequency of dyspeptic symptoms was not different between women with and without gallstones.[41] Misclassification could have occurred in this study through the inclusion of asymptomatic cases of gallstones among the non-case group and the inaccuracies inherent in dietary interview.[42,43] In both cases the influence would be toward a lack of association between the dietary factors studied and gallstone disease.

Our most important findings were the independent roles of fasting and dieting in the development of gallstones. We also found associations with energy intake and dietary fiber. Methodological constraints did not permit absolute rejection of the hypothesis that other nutrients affect gallstone disease hospitalization. □

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Cervical Cancer Screening: Who Is Not Screened and Why?

Linda C. Harlan, PhD, Amy B. Bernstein, ScD, and Larry G. Kessler, ScD

ABSTRACT

Background: The decline in death rates from cervical cancer in the United States has been widely attributed to the use of Papanicolaou (Pap) smears for early detection of cervical cancer.

Methods: Pap smear screening rates, beliefs about appropriate screening intervals and factors affecting screening were examined using 1987 National Health Interview Survey data.

Results: Results indicate that through age 69, Blacks are screened at similar or higher rates than Whites. Hispanics, particularly those speaking only or mostly Spanish, are least likely to have received a Pap smear within the last three years. Of women who had never heard of or never had a Pap smear, nearly 80 percent reported contact with a medical practitioner in the past two years, while more than 90 percent reported a contact in the past five years. Overall, the most frequently reported reason for not having a recent Pap smear was procrastinating or not believing it was necessary.

Conclusions: Thus, in developing screening programs, Hispanics, particularly Spanish speakers, must be targeted. In addition, educational programs should target unscreened women who forego the test due to underestimating its importance, procrastination, or because their medical care provider did not suggest the procedure. Women must be intensively educated that Pap smears should be scheduled routinely to detect asymptomatic cervical cancer. **(*Am J Public Health.* 1991;81:885–891)**

Introduction

Although cervical cancer is a readily curable disease if diagnosed and treated early, survival rates have not continued to improve over the mid-1970s and early 1980s.[1] Although these static survival rates may be related to better survivors being diagnosed in the *in situ* stage while women with poorer survival potential are diagnosed later, improved screening and early detection could increase survival rates and lower mortality. In addition, while overall rates of Pap smear screening are reasonably high, a significant number of women are not being screened.

Using the 1973 and 1985 National Health Interview Survey (NHIS), Makuc, *et al*, described changes in use of Pap smears by United States women.[2] They reported that poor women were less likely to use preventive services than nonpoor. They also found that between 1973 and 1985 Black women had substantially increased their use of Pap smears with rates exceeding those of White women. Older women also increased their screening rates, but their rates continued to be lower than younger women.

Using data from a telephone survey, Hayward, *et al*, found, in a multiple regression analysis, age was inversely related to the use of Pap smears.[3] Having health insurance, a higher level of education, being employed, and Black race (compared to Whites and Hispanics combined) were positively related to the use of Pap smears.

In a study conducted by 15 state health departments and the District of Columbia, Parker reported that 72 percent of the female participants had a Pap smear within the last year.[4] Rates decreased with increasing age, were higher in Blacks than in non-Hispanic Whites, and rose with increasing income. In this study, only 1 percent of the women had never had a Pap smear and 0.2 percent had never heard of the test. The knowledge and previous screening rates are much higher than other published rates.[2,3]

In a case-control study of cervical cancer patients, Celentano found that cases were significantly less likely than controls to have ever had a Pap smear or to have had regular Pap smears.[5] Cases were also older at the first Pap smear and more frequently reported not being informed that they should have routine Pap smears.

In England, a general medical practice studied women overdue for Pap smears.[6] Of the 118 women personally contacted, 47 percent said they had no major objection to the test, but it was of low priority. Another 24 percent gave reasons for not having the procedure which were based on incorrect information and 29 percent had strong objections to the test and would probably never consent to having a Pap smear.

Although research suggests Pap smear screening is relatively common, there are subgroups which remain resistant to screening efforts. The literature does not detail the underlying reasons for this failure. The 1987 National Health Interview Survey (NHIS) data offers an opportunity to explore these reasons. This paper will identify subgroups failing to utilize appropriate screening, as well as their underlying rationale.

From the National Cancer Institute. Address reprint requests to Linda C. Harlan, PhD, National Cancer Institute, Executive Plaza North, Room 313, 9000 Rockville Pike, Bethesda, MD 20892. This paper, submitted to the Journal May 31, 1990, was revised and accepted for publication January 8, 1991.

TABLE 1—Unweighted Number and Weighted Percent ( ) of Women Responding to 1987 NHIS Cancer Control Supplement by Race and Age

Age (years)	White	Black	Hispanic	Other	
18–29	2223 (24.3)	544 (31.3)	295 (37.4)	98 (37.9)	
30–39	2086 (21.2)	467 (22.0)	214 (22.0)	83 (26.5)	
40–49	1374 (15.6)	279 (16.6)	124 (15.3)	29 (11.0)	
50–59	1126 (12.9)	230 (12.5)	83 (11.6)	26 (13.3)	
60–69	1346 (12.8)	199 (9.4)	81 (8.5)	17 (7.1)	
70+	1678 (13.1)	199 (8.1)	51 (5.2)	16 (4.2)	
Total	9833 (99.9)	1918 (99.9)	848 (100)	269 (100)	12868

TABLE 2—Proportion of Women Never Having Heard of a Pap Smear and Proportion Compliant with Screening by Age and Race

	White				Black				Hispanic			
	Never Heard		Compliant		Never Heard		Compliant		Never Heard		Compliant	
Age (years)	$\bar{X}$	S.E.	$\bar{X}$	S.E.	$\bar{X}$	S.E.	$\bar{X}$	S.E.	$\bar{X}$	S.E.	$\bar{X}$	S.E.
18–29	2.6	(0.5)	83.3	(1.1)	3.4	(1.0)	85.3	(2.4)	16.0	(2.1)	67.2	(4.0)
30–39	0.6	(0.2)	85.7	(0.8)	1.4	(0.5)	90.8	(1.5)	10.0	(1.8)	76.9	(3.0)
40–49	2.0	(0.5)	74.9	(1.3)	0.5	(0.3)	89.1	(2.3)	10.3	(2.6)	75.5	(3.8)
50–59	0.6	(0.2)	72.0	(1.5)	4.5	(1.5)	70.3	(3.4)	19.7	(5.1)	53.6	(6.5)
60–69	1.7	(0.4)	61.6	(1.6)	5.1	(2.0)	61.1	(4.4)	19.4	(5.2)	43.4	(8.6)
70+	5.9	(0.7)	44.9	(1.4)	19.7	(3.6)	34.0	(3.8)	26.8	(6.4)	29.7	(6.2)
Total	2.1	(0.2)	73.3	(0.5)	4.1	(0.6)	78.9	(1.2)	15.1	(1.1)	65.0	(2.3)

Methods

Data from the 1987 NHIS were used to examine factors associated with the use of Pap smears by women in the United States. NHIS is a nationwide survey conducted by the National Center for Health Statistics (NCHS). When appropriate statistical methods are applied, estimates are representative of the US noninstitutionalized, civilian population. The 1987 sample design provided for an oversampling of Blacks and the Hispanic population directed at obtaining better estimates for these populations.

In addition to the core questions administered by NCHS, two supplemental questionnaires were added for the National Cancer Institute (NCI). One questionnaire focused on epidemiologic aspects of cancer, such as dietary intake. The second questionnaire addressed issues on cancer control, attitudes, and use of cancer screening. One of the two questionnaires was administered to one randomly selected adult member (aged 18 or over) of each of the approximately 44,000 households; 22,043 individuals completed the Cancer Control Supplement.

The NHIS Cancer Control Supplement included questions about the use of several cancer screening tests. With respect to Pap smears, women were asked, "Have you ever heard of a Pap smear test?" If a woman had not, she was not questioned further. Those women having knowledge of Pap smears were asked if they had ever had a Pap smear and when they last had one. Additional questions were asked about use for health problems, test results, the primary reason for not having the test, method used to advise them of results, and menstrual status, including hysterectomies.

Compliance was defined as having had a Pap smear in the past three years. The NCI's Working Guidelines suggest that after three or more consecutive normal annual Pap smears, the test may be performed less frequently at the discretion of the physician. Three years, often recommended as the appropriate interval, was chosen for the purpose of these analyses.[7]

Reasons for non-compliance were categorized into seven groups: 1) put off (the woman thought the Pap smear was unnecessary; she was simply putting it off or she had no problems and therefore did not have the test); 2) hysterectomy (the woman states she had a hysterectomy and therefore did not have a Pap smear); 3) cost (includes cost and having no insurance); 4) no MD (had no physician); 5) not recommended (the health care provider did not suggest a Pap smear or had said it was unnecessary); 6) embarrassment or fear on the part of the woman, or 7) other or unknown.

Totals, age-specific means and age-adjusted means were calculated using the SESUDAAN procedure.[8] This procedure takes into account the complex multistage sampling of the NHIS study. The RTI-LOGIT program was used for multi-variate logistic regression analysis taking into account the sample design.[9]

Results

The NHIS Cancer Control questionnaire was completed by 12,868 women age 18 and older. The unweighted distribution and weighted (to 1987 U.S. Census estimates) percentage of women by age and race can be found in Table 1. Women, other than non-Hispanic Blacks, non-Hispanic Whites or Hispanics, were excluded from further analyses because of the small sample size (n = 269). In addition, 3.6 percent of the women were unable to recall the date of their last Pap smear. These women were excluded from the analyses.

Age, Race, and Ethnicity

When the proportions of women never having heard of Pap smear are compared by age and race, the results are rather dramatic (Table 2). Many more Hispanic women have never heard of a Pap smear than either Black or White women. The proportion of women under the age of 70 having never heard of a Pap smear was 2.6 percent for White, 5.1 percent for Black, and 19.7 percent for Hispanic women.

As expected, compliance with screening is lower in Hispanics than for either Blacks or Whites (Table 2). Through age 29, compliance rates are similar for Blacks and Whites. From age 30 through 49, however, Blacks are significantly more compliant. In women age 70 and above, compliance rates for White

TABLE 3—Age-Adjusted Percent of Women Never Having Heard of or Interval Since Last Pap Smear by Primary Language Spoken

Language Spoken	Never Heard	Never Had	Last Pap Smear	
			<3 yrs	>3 yrs
Mostly Spanish	34.3 (5.0)	8.7 (2.6)	48.0 (6.5)	8.9 (2.8)
Both Equally	7.3 (3.6)	11.2 (3.7)	60.7 (5.0)	20.8 (4.2)
Mostly English	6.9 (2.7)	8.2 (2.5)	72.4 (3.5)	12.5 (2.9)

TABLE 4—Age-Adjusted Percent of Women Never Having Heard of or Interval Since Last Pap Smears by Poverty Index, Education, and Usual Source of Care

Variable	Never Heard	Never Had	Last Pap Smear	
			<3 yrs	>3 yrs
Poverty Index Ratio				
Below	6.2 (0.8)	10.9 (0.9)	64.8 (1.4)	18.1 (1.0)
At or above	2.5 (0.2)	6.5 (0.4)	75.2 (0.5)	15.9 (0.4)
Education				
<12 years	7.5 (0.6)	9.2 (0.8)	63.9 (1.2)	19.5 (0.8)
12 years	2.4 (0.3)	6.3 (0.5)	74.5 (0.8)	16.8 (0.6)
>12 years	1.5 (0.2)	6.2 (0.5)	79.0 (0.8)	13.3 (0.7)
Usual Source of Care				
HMO	1.1 (0.5)	4.7 (1.1)	85.2 (1.6)	8.9 (1.1)
Doctor's office	2.5 (0.2)	6.9 (0.4)	75.1 (0.6)	15.5 (0.4)
All other sources	4.7 (0.7)	6.5 (0.9)	73.7 (1.3)	15.1 (1.0)
No usual source	7.3 (0.9)	11.6 (1.0)	58.2 (1.4)	23.0 (1.3)
Total	3.3 (0.2)	7.2 (0.3)	73.3 (0.5)	16.1 (0.4)

women increase significantly over those for Black women. Through age 39, Hispanic women are less compliant than either Blacks or Whites. While above this age the rates vary greatly, the much lower rates for Hispanics suggest important differences in screening patterns.

To identify possible explanations for the low Hispanic compliance rate, this group was evaluated according to primary language: 1) only or mostly English (30.1 percent); 2) only or mostly Spanish (34.5 percent); or 3) both languages equally (32.2 percent) (undetermined 3.1 percent, excluded) (Table 3). When age-adjusted means were compared, nearly five times more Spanish speaking women had never heard of Pap smears as compared to English speakers. Much lower compliance rates for Spanish speakers were also apparent.

When compared by birthplace (US, Mexico, Cuba, and Puerto Rico or elsewhere), language appears to be a better predictor of screening compliance (data not shown). Using the self-categorization of Hispanic ethnic origin, Cubans were much less likely to be compliant than Puerto Ricans, Mexicans or "Other" Hispanics, who had similar rates of compliance. However, after adjusting for age, only "Other" Hispanics had significantly lower rates of compliance when compared to Mexicans. This may be important since "Other" Hispanics probably identifies women from South and Central America.

Other Variables

Rates of never having heard of Pap smear are over twice as high for women below the poverty index ratio, a measure of family income adjusted for household size (Table 4). When women above and below the poverty index are compared, it is apparent that the latter are less likely to have had a Pap smear in the past three years.

Education, as well as income, influences health behavior. Women with less than a high school education were three times as likely to have never heard of a Pap smear (Table 4). Women failing to complete high school were much more likely to be non-compliant with screening than those who graduated.

Menopausal status might also affect compliance rates. Women with non-surgical (primarily "natural") menopause were the least likely to be compliant. Premenopausal women were most likely to be screened in the last three years and women with surgical menopause (e.g. hysterectomies) fell between the two groups (data not shown). However, when rates are compared for these three groups after adjusting for age, no important differences in compliance are seen. Women who had undergone hysterectomies were included because excluding them did not appreciably alter other analytic results.

Attitudes toward cancer might influence use of screening Pap smears. When asked whether they could do very little to reduce their chances of getting cancer, women who agreed or strongly agreed were less likely to have had a Pap smear in the last three years than those who disagreed. Those who strongly disagreed were no more likely to have been screened in the last three years than women who disagreed. Women with "no opinion" were much less likely than any other category to have had a recent Pap smear.

For a variety of reasons—such as out-of-pocket expense, attitude of provider or setting where care is given—source of care might influence compliance with screening. Women in health maintenance organizations (HMOs) or other prepaid medical plans were more likely to have been screened within the last three years.* More than one-fourth of the women who reported doctor's office care were non-compliant. This is important since 70 percent of all women reported a physician's office as their primary source of care.

Timing of Pap Smears

Beliefs regarding proper screening intervals would also be expected to alter the frequency. When asked to describe appropriate screening intervals, responses ranged from "never" or "when problems occur" to "more frequently than once per year." Women who were unfamiliar with the procedure were not queried. Of those familiar, 77.9 percent described appropriate intervals as 13 to 36 months (one to three years) while 10.2 percent said annually (data not shown).

Whites were more likely than others to suggest a one- to three-year screening interval. However, Blacks and Hispanics were more likely to suggest a more frequent than annual screening. No important differences were identified among Hispanics when reported timing was examined by language spoken.

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*Bernstein AB, Thompson GB, Harlan LC: Differences in rates of cancer screening by usual source of medical care: Data from the 1987 National Health Interview Survey. (In press).

TABLE 5—Age-Adjusted Reported Reason for No Current Pap Smear by Race, Poverty Index, Education, Usual Source of Care, and Menopausal Status

	Put off		Hysterectomy		Cost		No MD		Not recommended		Embarrased Fear		Other/ Unknown	
	$\bar{X}$	S.E.	$\bar{X}$	S.E.	$\bar{X}$	S.E.	$\bar{X}$	S.E.	$\bar{X}$	S.E.	$\bar{X}$	S.E.	$\bar{X}$	S.E.
Age (years)														
18–29	50.8	(3.6)	0.3	(0.3)	5.2	(1.3)	5.0	(1.4)	11.3	(1.9)	7.0	(1.7)	20.4	(2.7)
30–39	47.0	(2.9)	7.8	(1.7)	9.9	(1.7)	8.8	(1.7)	1.0	(0.6)	6.6	(1.7)	18.9	(2.1)
40–49	45.4	(3.3)	15.6	(2.1)	5.6	(1.2)	7.7	(1.8)	5.9	(1.7)	1.9	(0.9)	17.9	(2.6)
50–59	42.2	(2.9)	20.5	(2.6)	6.0	(1.6)	5.7	(1.3)	8.8	(1.7)	1.2	(0.8)	15.5	(1.9)
60–69	50.2	(2.5)	14.1	(1.6)	3.2	(0.8)	6.6	(1.2)	10.8	(1.5)	2.3	(0.6)	12.9	(1.7)
70+	46.6	(2.0)	10.4	(1.2)	0.6	(0.3)	2.1	(0.5)	22.1	(1.8)	1.6	(0.4)	16.5	(1.8)
Race														
White	46.9	(1.5)	10.8	(0.8)	5.8	(0.7)	6.3	(0.7)	9.0	(0.9)	3.8	(0.7)	17.4	(1.0)
Black	49.9	(3.6)	2.8	(0.9)	6.1	(1.7)	6.1	(2.2)	13.1	(2.6)	2.5	(1.1)	19.4	(2.7)
Hispanic	47.8	(4.3)	2.5	(1.6)	4.0	(1.2)	4.2	(2.4)	11.5	(2.6)	8.7	(3.0)	21.4	(3.8)
Poverty Index Ratio														
Below	41.0	(2.9)	6.9	(1.9)	11.3	(2.3)	6.6	(1.7)	11.7	(2.0)	3.7	(1.4)	18.9	(2.4)
At/Above	48.8	(1.5)	10.2	(0.7)	4.3	(0.6)	6.5	(0.9)	8.5	(0.9)	4.5	(0.8)	17.3	(1.1)
Education (years)														
<12	45.6	(2.3)	8.6	(1.0)	7.7	(1.5)	4.7	(1.1)	8.5	(1.3)	6.6	(1.9)	18.3	(2.3)
12	49.0	(2.0)	10.5	(1.1)	6.1	(1.2)	5.5	(0.7)	9.6	(1.3)	3.9	(0.9)	15.4	(1.4)
>12	46.7	(2.6)	10.2	(1.2)	3.4	(0.7)	7.8	(1.6)	8.2	(1.1)	2.7	(0.8)	20.9	(1.9)
Usual source of care														
Office	47.6	(1.6)	11.4	(0.9)	5.2	(0.7)	4.8	(0.8)	9.7	(1.0)	3.7	(0.8)	17.6	(1.2)
HMO	42.5	(7.0)	8.8	(3.5)	3.7	(2.1)	3.5	(2.5)	12.3	(5.9)	1.3	(1.3)	27.9	(7.2)
All other	46.8	(3.0)	6.4	(1.3)	6.3	(1.8)	6.6	(1.6)	7.9	(1.6)	4.6	(1.5)	21.4	(2.4)
No USC	49.2	(3.2)	5.1	(1.2)	7.0	(1.5)	11.9	(1.9)	6.5	(1.7)	5.3	(1.6)	15.0	(2.0)
Menopausal Status														
Pre	51.3	(4.2)	—		7.5	(1.8)	5.8	(1.0)	5.7	(1.6)	3.8	(0.7)	20.0	(3.6)
Surgical	37.7	(2.7)	4.3	(1.5)	2.9	(0.9)	2.3	(1.3)	12.8	(1.6)	1.5	(0.9)	38.4	(1.5)
Post	52.2	(7.6)	—		2.3	(0.9)	10.0	(5.6)	11.9	(5.9)	9.1	(6.2)	14.4	(4.1)
Total	47.5	(1.4)	9.7	(0.7)	5.6	(0.6)	6.2	(0.7)	9.2	(0.8)	4.2	(0.6)	17.8	(0.9)

Reasons Reported for No Screening

Among those familiar with but non-compliant, the primary reason for not having a Pap smear was examined. The largest category of reasons given by women for being non-compliant was they believed it was unnecessary, had no problems, or had been procrastinating (Table 5). This category accounted for nearly half of the reasons given.

Reasons for not having a Pap smear varied greatly by age. Compared to other age groups, women in the youngest and oldest two groups more frequently stated that the physician had not recommended the test. The two youngest age groups were more likely than others to suggest embarrassment or fear as the reason for non-compliance.

By race no important differences were noted after adjusting for age, except that Whites were more likely than Blacks and Hispanics to suggest that the reason for no current Pap smear was having had a hysterectomy. When the age-adjusted reasons for lack of compliance were compared, surgical postmenopausal women gave "not recommended by the physician" more frequently and "procrastination" less frequently than premenopausal women (Table 5). Cost was a much more important factor for premenopausal women.

Logistic Regression Analyses

To determine factors associated with having heard of a Pap smear or compliance with screening, separate analyses using a backwards, stepwise logistic regression were performed for each as dependent variables (Table 6). Independent variables included: age (continuous), age squared (continuous), marital status (married vs divorced/separated, widow, never married), primary source of care (doctors' office vs HMO, all other sources, and no usual source), place of residence (urban vs rural), education (<12 years vs 12 years, >12 years), self-reported health status, menopausal status, attitude/knowledge about risk reduction (disagree vs strongly agree, agree, strongly disagree, no opinion) and language spoken (English vs Spanish, both languages equally). These variables were selected based on review of literature and/or previous univariate analyses.

In the logistic regression model, age and age squared were important predictors of knowledge of Pap smears, indicating a non-linear relationship of age and knowledge. Women who spoke only or mostly Spanish were less likely to have ever heard of Pap smears than women who spoke only or mostly English. No differences were seen between women whose usual source of care was a physician's office and those with any other usual source of care. However, women with no usual source of care were less likely to have heard of Pap smears than women using doctors' offices for care. Women with less than 12 years of education and those from urban areas were less likely to have ever heard of Pap smears, as were those who strongly agreed, agreed, or had no opinion about the risk statement.

A multiple regression analysis with compliance with screening (no, yes) as the dependent variable was performed (Table 7). These models included all independent variables listed above. Age, although an important predictor of compliance, has an inverted U-shape association with compliance. Compliance appears to peak around age 50, decreasing beyond that age 50 to a point lower than in the youngest age group. Speaking Spanish as the primary language was associated with a lower compliance rate. Never married women were less likely than married women to be compliant. Women who stated their primary source of care as an HMO were more likely to be compliant and women with no source of care were less likely to be compliant than women using a doctor's office. Women with 12 or more years of education were more likely to be compliant.

TABLE 6—Beta Coefficients and Standard Errors for Predictors of Never Heard of Pap Smears

Independent Variables	Beta	STD Error
Langage		
Mostly Spanish speaking	2.2399*	0.2838
Both languages	−0.1058	0.5340
Age	−0.1260*	0.0195
Age^2	0.0012*	0.0002
Usual Source of Care		
HMO	0.1166	0.4498
All other	0.4743*	0.2286
No care	1.1200*	0.1669
Education		
12 years	−0.7891*	0.1438
>12 years	−0.9945*	0.1659
Attitude		
Strongly agree	0.5432	0.3311
Agree	0.7394*	0.1679
Strongly disagree	0.3771	0.2589
No opinion	1.7424*	0.1385
Urban	0.4319*	0.1782

*Significant at $p < 0.05$.

TABLE 7—Beta Coefficients and Standard Errors for Predictors of Having Had a Pap Smear within the Past 3 Years

Independent Variables	Beta	STD Error
Langage		
Mostly Spanish speaking	−0.9981*	0.3459
Both languages	−0.4605	0.2447
Age	0.0034	0.0080
Age^2	−0.0004*	0.0001
Marital Status		
Divorced/Separated	−0.0567	0.0684
Widow	−0.1123	0.0714
Never Married	−0.9366*	0.0721
Usual Source of Care		
HMO	0.3724*	0.1157
All other	0.0457*	0.0953
No care	−0.7428*	0.0748
Education		
12 years	0.3582*	0.0612
>12 years	0.6381*	0.0739
Attitude		
Strongly agree	−0.0353	0.1737
Agree	−0.1859*	0.0694
Strongly disagree	−0.1657*	0.0757
No opinion	−0.5627*	0.0672

*Significant at $p < 0.05$.

Discussion

Data here confirm the findings of Makuc, *et al*, that screening in Black women now exceeds screening in Whites.[2] These data also confirm that screening decreases with increasing age and increases with education.[2,3] Although the screening rates are higher for Blacks, the mortality rate is lower for Whites.[1] This may be due to the fact that only recently have screening rates increased above those for Whites. It is also possible that follow-up of positive Pap smears is not as complete for Black women.

Data suggest that Hispanic women are at an increased risk of mortality and morbidity from cervical cancer.* Despite this fact, it is clear that these women are not being adequately screened. Several factors appear to contribute to this lack of screening.

One might suspect that older Hispanic women would be less inclined to learn English and age would therefore confound the relationship between language spoken and compliance. Even after adjusting for age and other variables, women speaking only or mostly Spanish remained less likely to have been screened. Language spoken may be a surrogate measure for acculturation. There may be a cultural avoidance of medical tests such as Pap smears. Although the estimates were somewhat unstable because of small numbers, many more Spanish speakers (17.7 percent) than English speakers (1.5 percent) reported the primary reason for no current Pap smear was because of embarrassment or fear. However, over 11 percent (age-adjusted) of the Spanish speaking women reported they were non-compliant with screening because their physician had not suggested it or had said it was unnecessary. Also contributing may be the fact that few medical care providers are able to fully communicate in Spanish and therefore are unable to convey the importance and need for a Pap smear.

Non-compliant Hispanic women account for approximately 1.6 million unscreened women and Hispanics are the fastest growing segment of our population.[10] Any program to increase screening participation will ultimately need to address existing cultural differences, especially for Spanish speakers.

Usual source of medical care appears also to influence compliance. Women whose usual source of care was an HMO were more likely to be screened than women whose usual source of care was a physician's office. This is an important source of unscreened women since most women still have their primary source of care in a physician's office. Nationwide this would account for about 18.6 million women who were unscreened in the last three years. Some physicians who are the primary source of care for a woman may not provide Pap smear tests, e.g. the internist may think gynecologists should do the test, and a woman may not have regular gynecological care (data not available in this data set). However, screening would probably increase if these physicians provided referrals.

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*Horm JW: Cancer Incidence, Survival, and Mortality Among Racial and Ethnic Groups in the U.S. (unpublished).

Screening compliance might be affected by the manner in which the results are communicated. Makuc, *et al*, found that the rates for Pap smears in Blacks had increased substantially between 1974 and 1985, suggesting a targeted effort was made in this group.[3] As part of that effort of reinforcing compliance, physicians might be giving test results in person. Indeed, more than half of the Black women received their Pap smear results in this way. Women with less education and lower incomes were more likely to receive the results of their Pap smears in person; however, these women were less likely to have had a Pap smear within the last three years. In addition, women whose usual source of care was an HMO have the highest rate of Pap smear screening, but most commonly these women are not contacted with normal results. Even with diagnostic Pap smears, HMO-enrolled women were still more likely not to be contacted with normal results. These findings suggest that in-person results may instead reflect the provider concern that follow-up might be difficult when results are abnormal, but does not support the hypothesis that in-person results reinforce screening.

Knowledge of proper screening intervals would be expected to influence compliance with screening. When queried about the proper frequency of screening, 75 percent of the women believed the appropriate screening interval was one to three years. However, when examined by race, more Black and Hispanic women believed screening should be performed with a more than annual frequency. Despite the fact that Black women more often reported an inappropriate frequency for Pap smears, they were still more likely to be screened than White women. In addition, while Hispanic women were more likely to say that Pap smears should be performed more than annually, they were the least likely to have received them at the appropriate rate. This suggests components other than knowledge may be encouraging the screening of Black women, since knowledge does not necessarily translate into proper screening practice.

Reasons varied for non-compliance, but nearly 60 percent of the reasons basically revealed a lack of appreciation of the importance of screening. An estimated 12 million US women could be included in this group along with 2.5 million who have never heard of Pap smears. These findings suggest that this population is not adequately educated about the importance of routine screening, lack encouragement from medical providers, or have substantial misunderstanding about the messages from them regarding Pap smears.

Women may be getting mixed messages about screening, with the recommendations for screening frequency seemingly unclear. In addition, the controversy regarding the accuracy of the reading of Pap smears, which has been well reported in the lay press, might convey a message that screening is unimportant or results are of questionable quality.[11] A clear, unambiguous statement about the benefits and timing of Pap smears needs to be emphasized.

Lack of screening compliance can be attributed to multiple causes. Any effective screening intervention must include a multi-dimensional strategy that will successfully address these variables and influence behavior to generate better compliance among women at risk of developing cervical cancer. □

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Public Health Briefs

Sustained Effects of an Educational Program to Reduce Sales of Cigarettes to Minors

David G. Altman, PhD, Lolly Rasenick-Douss, Valodi Foster, MPH, and Joe B. Tye, MBA

ABSTRACT

We report 1-year follow-up data from a sample of stores participating in a 6-month community-wide educational effort to reduce cigarette sales to minors in Santa Clara County, California. The proportion of over-the-counter sales to minors at the 1-year follow-up illustrated that although statistically significant reductions were maintained 6 months after the intervention ended, recidivism occurred. Suggestions for achieving long-term reductions in sales to minors are offered. (*Am J Public Health*. 1991;81:891–893)

Introduction

Although there are many factors that influence the decisions that adolescents make to initiate or continue smoking, one that has received only recent attention is access to tobacco. Few smokers begin smoking after their teenage years, and the earlier a young person begins using tobacco, the less likely it is that he or she will be able to quit later.[1] Access to tobacco is conducive to developing and maintaining a tobacco addiction. In field trials around the country, minors' attempts to purchase tobacco from stores and vending machines have been successful more than 50% of the time.[2–10] A survey of 10th graders in two Minnesota communities found a wide diversity of locations from which tobacco had been purchased.[11] In the United States, 947 million packs of cigarettes worth $1.23 billion are sold illegally to minors each year, and the tobacco industry derives 3% of its profits ($221 million) from such sales.[12]

Although most states have laws regulating sales to minors, these laws are rarely enforced.[5,7,13–16] An analysis of 11 active state and local enforcement efforts and 1 200 interviews with health experts, students, parents, and vendors from 18 states found that although 44 states and the District of Columbia have laws prohibiting the sale of tobacco to minors, only 32 violations in 5 states were documented.[13]

Previously, we reported the 6-month effects of a comprehensive merchant and community education program to reduce cigarette sales to minors.[2] One-year follow-up data are reported here.

Methods

In January 1988, 412 diverse stores in Santa Clara County, California that sold cigarettes over the counter and 30 outlets that had cigarette vending machines were visited by minors aged 14 to 16 years. In July 1988, after several months of intervention, which included comprehensive community education, direct merchant education, and contact with chief executive officers of chains and franchises, all stores were revisited by minors. Although the interventions concluded in July 1988, we wanted to assess their sustained effects. The primary data reported here were collected in January 1989, 1 year after the posttest and more than 6 months after the intervention ended. At this follow-up, a random sample of 25% of stores that sold tobacco over the counter, stratified by city (n = 7) and type of store (chain, franchise, independent), was selected and those that were still in business were visited by minors (n = 97). Because a new state law enacted in January 1989 made it illegal for minors to purchase tobacco, we obtained letters from local law enforcement officials allowing teenaged boys and girls to conduct this follow-up without being arrested.

Results

Statistically significant reductions from pretest to follow-up were generally

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David G. Altman and Lolly Rasenick-Douss are with the Center for Research in Disease Prevention at Stanford University School of Medicine, Valodi Foster is with Health Works Associates, and Joe B. Tye is with Baystate Medical Center.

Requests for reprints should be sent to Dr. David G. Altman, Center for Research in Disease Prevention, Stanford University School of Medicine, 1000 Welch Road, Palo Alto, CA 94304-1885.

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TABLE 1—Cigarette Sales to Minors

	Pretest (all stores)	Pretest (stores with pretest and 1-year data)	6-Month Posttest (all stores)	1-Year Follow-up	P Value	% Change Pretest to 1-Year Follow-up and 95% CI[a,b]
All stores	74% (412)[c]	76% (97)	39% (408)	59% (97)	.01	−18% (−4%, −31%)
Chain stores	62% (174)	70% (40)	27% (174)	58% (40)	NS[d]	−13% (11%, −36%)
Franchise stores	81% (110)	82% (28)	50% (110)	50% (28)	.02	−32% (−5%, −59%)
Independent stores	85% (121)	79% (29)	47% (121)	69% (29)	NS	−10% (10%, −31%)
Gas stations	84% (50)	91% (11)	55% (50)	55% (11)	.10	−36% (7%, −80%)
Liquor stores	83% (103)	87% (23)	46% (103)	65% (23)	.06	−22% (1%, −44%)
Convenience stores	81% (103)	82% (28)	48% (103)	54% (28)	.03	−29% (−2%, −55%)
Pharmacies	60% (67)	53% (15)	28% (67)	53% (15)	NS	0% (−37%, 37%)
Grocery stores	60% (89)	65% (20)	22% (89)	65% (20)	NS	0% (−34%, 34%)
Sale to male minor	71% (180)	80% (70)	35% (87)	53% (70)	<.001	−27% (−11%, −43%)
Sale to female minor	77% (122)	65% (26)	47% (73)	77% (26)	NS	12% (−16%, 39%)
Warning sign posted	3% (13)	1% (97)	23% (92)	25% (97)	<.001	24% (14%, 34%)
Minor asked age	24% (100)	21% (96)	48% (196)	34% (96)	.05	14% (0%, 27%)

[a] % change equal to % in pre-test minus % in 1-year follow-up (% rounded).
[b] 95% confidence interval was calculated for change by using sample with data at pretest and 1-year follow-up.
[c] The number of stores for each variable is provided in parentheses.
[d] NS indicates a significance level greater than .10.

obtained even though sales increased between posttest and follow-up. Table 1 presents pretest, 6-month posttest, and 1-year follow-up data. Pretest data are provided for the entire sample of stores and for the sample of stores visited at both pretest and 1-year follow-up. The McNemar nonparametric test for the significance of changes and confidence intervals were calculated on the difference in percent change between pretest and 1-year follow-up for stores with both pretest and 1-year follow-up data.[17]

Discussion

There are discrepancies between what boys and girls are taught in school-based prevention programs and what they observe and experience outside of the classroom. In the classroom, they are told not to smoke, yet in the community, adolescents observe an abundant supply of tobacco, a willingness on the part of merchants to sell them tobacco, a lack of enforcement of laws regulating sales to minors, and very little evidence that society views the problem as a priority. Changing this discrepancy should be a public health priority.

The data reported here illustrate that a comprehensive merchant and community education campaign can achieve sustained effects in reducing sales to minors. However, without continued intervention, monitoring, and we now believe, sanctions, recidivism will occur. Because we wanted to test explicitly the effectiveness of a purely educational intervention, we did not encourage law enforcement to cite stores that sold tobacco to minors. Follow-up data reported here as well as data from other projects that utilized the sanctions of enforcement convince us of the importance of combining education of merchants and the public with the bite of enforcement.[18,19] The experience of the Solano County Cancer Prevention Program in California[19] suggests, that communities will respond more favorably to punitive enforcement efforts if educational interventions are implemented first. Once nonpunitive educational interventions relying on merchants' voluntary compliance are implemented and illegal sales are not reduced entirely, enforcement becomes a more acceptable option. Preceding enforcement with education also allows merchants an opportunity to change their business practices without the immediate threat of punishment.

It is reasonable to hypothesize that if minors have a difficult time obtaining tobacco, then they may be prevented from experimenting with and later becoming addicted to it. Along these lines, studies of school smoking policies have found that making smoking inconvenient may reduce tobacco use by adolescents.[20–22] Without doubt, there will always be some tobacco use by adolescents, just as there is illegal use of marijuana, crack, and other hard drugs. But the more risk young people must take to obtain tobacco and the more societal norms reject tobacco use by young people, the greater the likelihood that young people will not use tobacco. It is also reasonable to ask whether interventions to reduce tobacco sales to adolescents will actually reduce access or simply change where access occurs. Unless sales to adolescents are totally stopped, which is unrealistic in most communities, adolescents who really want tobacco will be able to get it. If, however, community norms and merchant practices about sales to adolescents are applied consistently and comprehensively, we predict that access will truly be reduced. As this study and others have shown, educating the community, merchants who sell tobacco, public policymakers, and law enforcement officials can contribute to reducing the sale of tobacco products to minors.

The 1990s will undoubtedly be a landmark decade for eliminating minors' access to tobacco now that communities around the country have begun addressing the problem. For example, in 1990, research in Minnesota resulted in national attention when dozens of communities throughout the state passed ordinances restricting minors' access to vending machines (Jean Forster, University of Minnesota, personal communication, 1990). The US government also has spoken out on the problem. In May 1990, HHS Secretary Sullivan recommended model state legislation to control minors' access.[14,15] These recommendations dealt with retail licensing, minimum age, graduated penalties for illegal sales, warning signs, enforcement, and vending machines.

Our research suggests that although long-term effects of comprehensive educational programs to reduce tobacco sales to minors can be achieved, interventions must be sustained and rely on multiple intervention strategies. Future intervention research should address minors' access to smokeless tobacco, challenges of intervening in densely populated urban areas, the effectiveness of education combined with enforcement and policy change, the sustainability of intervention effects, and the relationship between reducing tobacco access and youth tobacco use. ☐

Acknowledgments

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Hormone–PMS Relationship More Complex Than Previously Thought

Contrary to popular belief, a new study suggests that premenstrual syndrome (PMS) is not simply related to changes in the levels of reproductive hormones that occur during the premenstrual phase of a woman's cycle. Instead, PMS may be triggered by reproductive hormonal events, as yet unidentified, occurring before the premenstrual phase, or it may be the result of a cyclic mood disorder that is synchronized with (but not caused by) the menstrual cycle.

An estimated 5% of American women are affected by PMS and experience behavioral, emotional, and physical symptoms severe enough to impair their social or occupational functioning. Symptoms include sadness, irritability, anger, tension, loss of energy, and changes in appetite and sleeping patterns.

The study entitled "Lack of Effect of Induced Menses on Symptoms in Women with Premenstrual Syndrome," which appeared in the April 25 issue of the *New England Journal of Medicine*, was conducted by researchers at the National Institute of Mental Health (NIMH) and the National Institute of Child Health and Human Development, in Bethesda, Maryland. NIMH scientists Peter J. Schmidt, MD, and David R. Rubinow, MD, headed the investigation.

The study helps explain why hormone therapy attempting to correct a presumed abnormality during the premenstrual phase is not effective in treating PMS. The findings are consistent with results from previous studies showing the inefficacy of therapy with the hormone progesterone.

"This important study dramatically alters our understanding of and treatment approaches for PMS. It provides valuable new information about the disorder and will stimulate further research to help the millions of women suffering from PMS." said Alan I. Leshner, PhD, NIMH acting director. "Scientifically, the fact that the hormone–PMS relationship is not as simple as previously believed makes PMS a useful model for understanding the very complex relationship between biology and behavior."

What Scientists Funded by the Tobacco Industry Believe About the Hazards of Cigarette Smoking

K. Michael Cummings, PhD, MPH, Russell Sciandra, MA, Amy Gingrass, and Ronald Davis, MD

ABSTRACT

Despite overwhelming evidence documenting the hazards of cigarette smoking, the tobacco industry denies that smoking has been proven to cause disease. The industry professes a desire to clear up the smoking and health "question" and often points to its support of the Council for Tobacco Research (CTR) as evidence of its interest in investigating the health dangers of smoking. This paper presents results of a survey of CTR-funded scientists regarding their beliefs about the health dangers posed by smoking cigarettes. The vast majority of scientists funded by the CTR believe that cigarette smoking is an addiction that causes a wide range of serious, often fatal, diseases. This result suggests that the tobacco industry is unwilling to accept even the opinions of scientists it has deemed worthy of funding. Scientists should consider the ethical implications of accepting funds from the CTR and other tobacco industry-supported institutions. (*Am J Public Health*. 1991;81:894–896)

Introduction

The US Surgeon General has stated that "smoking represents the most extensively documented cause of disease ever investigated in the history of biomedical research."[1] Despite overwhelming scientific evidence against cigarettes, the tobacco industry continues to assert that controversy, debate, and uncertainty exist among scientists concerning smoking as an important cause of illness.[2–4]

In 1972, a confidential memorandum from a Tobacco Institute (TI) vice-president described TI policy as "creating doubt about the health charge without denying it, advocating the public's right to smoke without actually urging them to take up the practice, and encouraging objective scientific research as the only way to resolve the question of the health hazard."[5] Industry spokespersons often point to the industry's support of the Council for Tobacco Research (CTR) as evidence of corporate interest in obtaining scientific evidence on the "alleged" relationship between tobacco use and disease.[3]

The CTR, formed in 1954 by cigarette manufacturers, describes its primary mission as support of research into questions of tobacco use and health.[6] The council awards peer-reviewed research grants to independent scientists who are assured complete scientific freedom in conducting and publishing their studies. Since 1954, the council has provided more than $150 million for 1,108 original studies by more than 700 scientists.[6] In 1989, the CTR listed 204 active projects.

We present results of a survey of CTR-funded investigators, which characterizes what the investigators believe regarding the health effects of tobacco.

Methods and Materials

The study population included the principal investigators of research projects funded by the CTR in 1989.[6] Of 204 investigators listed, 179 were located at universities or institutions in the United States. This survey was restricted to those working in the United States.

Eligible survey participants were mailed a one-page questionnaire assessing their beliefs about the relation between cigarette smoking and various health complications and asking them to rank the importance of 10 different areas of tobacco research. Respondents were asked about their current and past research on tobacco and their cigarette smoking status. A cover letter stated that we were surveying "scientists who had published studies on smoking and health and/or have received research support from organizations interested in the tobacco and health issue."

Of 179 questionnaires mailed to eligible scientists in July 1990, 13 were returned with an incorrect mailing address. A total of 77 completed questionnaires were returned, which represents a response rate of 46% (77/166). No further attempt was made to elicit response. A comparison of responders and nonresponders revealed no significant differences in academic credentials (PhD vs MD), institutional affiliation (university vs other), or the nature of the CTR-funded project (i.e., a tobacco study vs a nontobacco study).

Results

Respondents were asked to "indicate the degree to which you believe the scientific evidence suggests a causal relationship with cigarette smoking" for

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K. Michael Cummings, Russell Sciandra, and Amy Gingrass are with the Department of Cancer Control and Epidemiology, Roswell Park Cancer Institute, and Ronald Davis is with the Office on Smoking and Health, Center for Chronic Disease Prevention and Health Promotion, Centers for Disease Control.

Requests for reprints should be sent to K. Michael Cummings, PhD, MPH, Department of Cancer Control and Epidemiology, Roswell Park Cancer Institute, Elm and Carlton Streets, Buffalo, NY 14263.

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eight separate health complications. With the exception of bladder cancer, nearly all respondents rated the relationship between smoking and illness as "strong" or "moderate" (Table 1). Only 1 of 77 respondents was a current smoker. Nationally, approximately 29% of adults smoke.[7]

Only 22% of respondents indicated that any of their current or past research focused on the health effects of tobacco use.

Ninety-four percent of respondents agreed with the statement "smoke from someone else's cigarette is harmful to a nonsmoker"; 91% agreed that most deaths from lung cancer are caused by smoking; and 76 of 77 agreed with the statement "cigarette smoking is addictive" (Table 2).

Rankings of the importance of 10 areas of research on tobacco varied widely (Table 3). Overall, research on preventing tobacco use received the highest ranking, followed by smoking cessation methods.

TABLE 1—Beliefs About the Scientific Evidence Suggesting a Causal Relationship Between Cigarette Smoking and Various Illness Conditions

	Strength of Causal Relationship with Cigarette Smoking:			
Illness/Condition	Strong %	Moderate %	Slight %	Not Established %
Emphysema	88.9	9.7	-	1.4
Chronic bronchitis	77.1	20.0	1.4	1.4
Heart disease	67.1	27.4	2.7	2.7
Low birth weight	51.6	30.6	12.9	4.8
Lung cancer	93.2	5.4	1.4	—
Cancer of the mouth, throat	76.1	19.4	1.5	3.0
Bladder cancer	24.6	56.1	15.8	28.1
Shorter life expectancy	76.4	16.7	4.2	2.8

TABLE 2—Beliefs About the Dangers of Environmental Tobacco Smoke, Smoking as a Primary Cause of Lung Cancer, and Smoking as an Addiction

Statement	Agree Strongly %	Agree Somewhat %	Disagree Somewhat %	Disagree Strongly %
The smoke from someone else's cigarette is harmful to a nonsmoker	52.9	41.4	5.7	—
Most deaths from lung cancer are caused by cigarette smoking	66.2	25.4	8.5	—
Cigarette smoking is addictive	83.3	15.3	1.4	—

TABLE 3—Mean Rankings of the Importance of 10 Areas of Research on Tobacco

Research Area	Mean Rank*	Standard Deviation	Percent Ranking Area as Top Priority
Tobacco use prevention	5.4	2.9	19%
Smoking cessation methods	5.2	2.8	14%
Neoplastic diseases	5.1	3.1	21%
Cardiovascular diseases	5.1	2.4	8%
Pregnancy and infant health	4.9	2.3	7%
Nonneoplastic respiratory diseases	4.6	2.3	—
Passive smoking	4.5	2.7	9%
Chemistry, pharmacology, toxicology	3.5	2.8	6%
Tobacco economics and legislation	3.4	3.4	12%
Tobacco advertising	3.4	2.9	4%

*Ranks range from low = 1 to high = 10.

Discussion

The low response rate is not unusual for a mailed questionnaire survey,[8] but does potentially limit the generalizability of the findings. The comparison of respondent and nonrespondent characteristics does not suggest any systematic response bias. It is possible that among those not responding to the survey were some who feared retribution from the tobacco industry. Such fear, if it existed, would likely be among those whose views are congruent with the majority of the respondents. It is also possible that those with less scientifically acceptable positions or greater commitment to the tobacco industry were less likely to respond, though it would seem to be in the industry's interest to have its views represented in such a survey, if possible. For these reasons, we believe the results accurately represent views of scientists funded by the CTR.

The survey shows that most scientists funded by the CTR believe cigarette smoking is an addiction that causes a wide range of serious, often fatal, diseases. This finding conflicts with the tobacco industry's description of the scientific community as divided on the question and indicates the industry does not accept the opinions even of scientists whose research it funds. Although acknowledging a need for additional investigation of the mechanisms linking smoking and disease, respondents gave the highest priority to research on tobacco use prevention and cessation. None of CTR's active research projects relate to these topics. Despite its stated mission to fund research into the etiology of diseases "alleged" to be related to tobacco use, only 1 in 6 CTR-funded scientists reported conducting research focused on the health effects of tobacco.

We suggest that rather than sponsoring a genuine "program of research into questions of tobacco use and health"[6] the CTR is a public relations vehicle intended to foster a false impression that cigarette manufacturers are interested in investigating the smoking and health "question." We believe such misuse of science raises serious ethical questions for scientists who accept funding through CTR or similar industry-supported entities. Even assuming that adequate funding is not available elsewhere, tobacco industry-supported scientists must ask themselves whether the value of their research in expanding the body of biomedical knowledge outweighs its utility in furthering the corporate interests of a business which kills 434,000 Americans every year.[9] ☐

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Gun-Related Deaths Increase Further for Black and White Young Men

The Department of Health and Human Services (HHS) reported today that one out of five deaths of teens and young adults in 1988 was gun related. Among young Black males, close to half of all deaths were firearm related.

Researchers at HHS' National Center for Health Statistics said the firearm-related death rate among Black and White males aged 15 to 24 years had declined in the early 1980s, but then rapidly increased from 1984 to 1988, especially for teenagers and particularly among Black males. Between 1984 and 1988, the firearm death rate among teenagers increased by over 40%, rising 20% from 1987 to 1988 alone, and reaching its highest level to date, 17.7 deaths per 100,000.

"This study shows that for the first time, the firearm death rates for both White and Black male teenagers exceeded the mortality from all natural causes of death," said HHS Secretary Louis W. Sullivan, MD. "For young Black males in particular, the excessive firearm and homicide death rates are appalling and heart-rending.

For Black males aged 15 to 19 years, the firearm death rate and firearm homicide rate more than doubled from 1984 to 1988, with an increase of 38% from 1987 to 1988 alone. By contrast, the nonfirearm homicide rate remained relatively stable. Black male teenagers were almost three times as likely to die from gun-related deaths than from all natural causes of death. For White males 15 to 19 years, the firearm death rate in 1988 for the first time exceeded that of natural causes, by 11%.

HHS Assistant Secretary for Health James O. Mason, MD, who heads the US Public Health Service, said, "The statistics show an American epidemic—without parallel in any other industrialized nation of the world. Physicians and other citizens need to work together to devise the same kinds of educational and prevention techniques that we use to attack other epidemics."

The findings are contained in a study, "Firearm Mortality Among Children, Youth and Young Adults, 1979–1988," produced by NCHS, a part of the Centers for Disease Control. The study examines homicide, suicide, and unintentional firearm deaths among those aged 1 to 34 years.

Does Cigarette Smoking Have an Independent Effect on Coronary Heart Disease Incidence in the Elderly?

Richard Benfante, PhD, MPH, Dwayne Reed, MD, PhD, and John Frank, MD, MSc

ABSTRACT

In order to evaluate the effects of cigarette smoking on coronary heart disease (CHD) in elderly persons in the Honolulu Heart Program, 1,394 men between ages 65 and 74 were followed during an average 12-year period for new cases of nonfatal myocardial infarction and fatal CHD. Incidence rates increased progressively in individuals classified at baseline as never, former, and current smokers, respectively. The absolute excess risk associated with cigarette smoking was nearly twice as high in elderly compared with middle-aged men. (*Am J Public Health.* 1991;81:897–899)

Introduction

While cigarette smoking is well established as a major risk factor for coronary heart disease (CHD) in middle-aged populations,[1–3] the question of whether it is a risk factor for CHD in the elderly has recently arisen because some studies have reported a diminished effect of cigarette smoking on CHD in older persons.[4–9] This reported lack of effect of cigarette smoking on CHD in the elderly is puzzling because the deleterious effects of smoking on CHD risk have been implicated at various stages of atherosclerosis, a pathologic process that generally progresses with age.[9–12] We had an opportunity to examine this question in an elderly group of the Honolulu Heart Program (HHP) cohort. The present analysis was undertaken to determine whether cigarette smoking behavior in the elderly predicts subsequent disease to the same degree as in middle-aged men.

Methods

In the HHP cohort, 1,394 men between ages 65 and 74 and free of clinical cardiovascular disease were followed for 12 years after baseline examination for new cases of definite CHD (non-fatal myocardial infarction and fatal CHD). Details of the differential diagnosis of CHD, surveillance procedures, and case ascertainment in the HHP have been previously described.[13–16]

In order to examine the effect of smoking on CHD risk, age-adjusted incidence rates were calculated by smoking status (current, former, and never) at baseline and expressed in terms of person-years of follow-up. The independent role of cigarette smoking on CHD risk was evaluated using a Cox proportional hazards model that included other major risk factors for CHD.[17,18] Relative risks for cigarette smoking were based on the risk ratio between current smokers and never smokers.

In order to see whether the effect of cigarette smoking on CHD risk was different in elderly compared to middle-aged men, 3,429 of the youngest men in the cohort (ages 51–59 at the examination) were studied as well.

Results

Average values for selected smoking characteristics in elderly and middle-aged men are presented in Table 1. Current elderly smokers on average smoked 20 cigarettes per day compared to 24 cigarettes per day for middle-aged men. Current elderly smokers smoked an average of 10 years longer than middle-aged smokers. For both elderly and middle-aged men, former smokers had higher daily consumption rates than current smokers.

In both age groups there was a consistent and progressive increase in CHD incidence rates by increasing level of smoking experience (Figure 1). Never smokers showed the lowest rates of CHD, followed by former smokers, with current smokers exhibiting the highest. When the relative risks for elderly and middle-aged men were compared in a Cox multivariate lifetable regression that included other major risk factors for CHD (systolic blood pressure, serum cholesterol, diabetes history, body mass index, and alcohol intake) the relative risk for the elderly was 1.62 (95% CI = 1.01, 2.61) compared to 1.80 (95% CI = 1.25, 2.61) for the younger men (Table 2). The difference in relative risks (relative risk ratio) between elderly and middle-aged men was not statistically significant. When an age-smoking interaction term was included in the model, it was also not statistically significant. From the perspective of excess risk

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From the Honolulu Heart Program (Benfante, Frank) and the National Heart, Lung and Blood Institute (Reed). Address reprint requests to Richard Benfante, PhD, MPH, Honolulu Heart Program, Kuakini Medical Center, 347 N. Kuakini, Honolulu, HI 96817. This paper, submitted to the Journal May 8, 1990, was revised and accepted for publication August 15, 1990.

TABLE 1—Average Values for Smoking Characteristics in Elderly and Middle-Aged Smokers at Baseline Examination

	Smoking Status by Age Group at Examination			
	Ages 51–59		Ages 65–74	
Smoking Characteristics	Past N = 1167	Current N = 1303	Past N = 527	Current N = 406
Cigarettes/day	27.5	23.8	26.3	19.7
Years Smoked	24.3	36.1	31.2	46.5
Years Quit	11.6	—	14.9	—

TABLE 2—Relative Risks* for Coronary Heart Disease in Elderly and Middle-Aged Men Based on Current Smoking Status at Baseline Examination

Age Group	N	b	S.E.	RR†	95% CI
65–74	1394	0.485	0.243	1.62	1.01, 2.61
51–59	3429	0.589	0.188	1.80	1.25, 2.61

*Based on Cox multivariate regression (includes age, cigarette smoking, systolic blood pressure, diabetes history, serum cholesterol, body mass index and alcohol intake).

†Determined by taking the exponent of the b-coefficient for smoking (Current—Never) in multivariate setting.

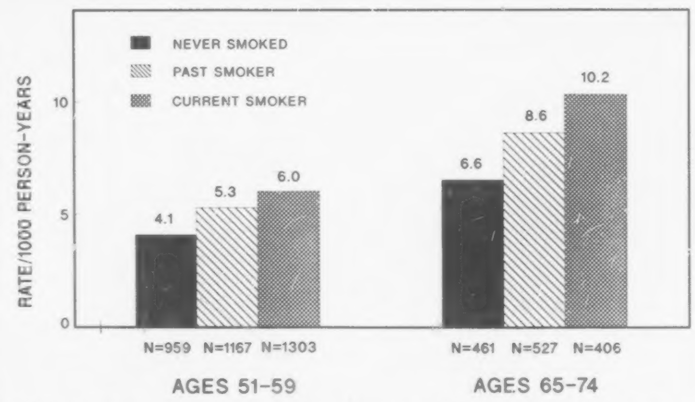

FIGURE 1—CHD Incidence by Smoking Status in Middle-Aged and Elderly Men

for CHD, the difference between current and never smokers for the elderly was 3.6 events per 1,000 person-years, compared to 1.9 events for middle-aged smokers.

Discussion

The results of the present study show that cigarette smoking continues to be an independent predictor of CHD incidence in this cohort of Japanese-American men over age 65, and that the effect is relatively undiminished compared with middle-aged men. CHD incidence for past smokers was intermediate between current and never smokers in both age groups. This is consistent with other evidence that the beneficial effects of quitting are effective later in life as well as during middle years.[1,19-21] It is also important to note that while the relative risk for cigarette smoking on CHD was similar in elderly and middle-aged men, there was nearly a two-fold increase in the excess absolute risk in the elderly men. Thus, since CHD rates increase exponentially with age, the public health impact of a risk factor such as smoking may be greater in older persons even though the relative risk is unchanged.

While a few other studies of older persons have reported an association of cigarette smoking with cardiovascular disease,[19-22] some prospective cohort studies report a diminished effect in the elderly.[4-9] Most notable is the Framingham study which has recently reported that the relationship between cigarette smoking and CHD disappears in both men and women over age 65.[7,8] Several factors may account for differences between the results of Framingham and Honolulu. One is the different analytical methods of relating risk factors to CHD. We used a single baseline measure with an average 12 years of follow-up, compared to the use of the most recent measure from a series of biennial examinations in Framingham. If there was a change in risk factor level, the most recent measure would be used; for example, if a smoker quit, he would be coded as a nonsmoker and then followed for subsequent disease.

Another factor is the potential intervention effect of numerous repeat examinations in Framingham versus just a few examinations in Honolulu. Subjects and their physicians could be notified of high risk levels numerous times in Framingham, but only a few times in Honolulu. Thus, behavioral change or medical intervention would be more likely to occur in Framingham.

There is also the selective attrition effects related to the different time periods of the two studies. The first 20 years of the Framingham follow-up were 1948–68, a period when CHD rates were the highest known in the US. The Honolulu study started in 1965 during a period of declining CHD mortality rates.[23,24] It is possible that these secular effects resulted in differences in the proportion of susceptible subjects reaching the elderly group in the two studies. In addition, the sample of men between ages 65–74 in Honolulu was about twice that for the same age group for men in Framingham, providing a distinct advantage in statistical power.

In conclusion, the present study supports the hypothesis that the biological or clinical effects of such a strong risk factor as cigarette smoking in middle age are maintained over time with increased exposure, and that the benefits of quitting accrue regardless of age. It is therefore recommended that the cessation of cigarette smoking be promoted as part of an overall strategy for prevention of CHD at all ages. ☐

Acknowledgments

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Staff Involvement and Special Follow-up Time Increase Physicians' Counseling About Smoking Cessation: A Controlled Trial

Carol Duncan, RD, MPH, Morton J. Stein, MD, and Steven R. Cummings, MD

ABSTRACT

We compared the counseling behaviors of two groups of health maintenance organization physicians: one group received training about smoking cessation counseling; the other group received the same training plus staff support and appointment time specially designated for follow-up of smokers. We interviewed patients after their office visits to measure smoking counseling. The group receiving staff support and designated follow-up time counseled more and made more follow-up appointments about smoking. (*Am J Public Health.* 1991;81:899–901)

Introduction

Only a minority of smokers recall being counseled to quit by physicians[1,2] and internists who do counsel smokers seldom use effective strategies such as setting quit dates and making follow-up appointments.[3,4] We have found that even after intensive training in counseling, physicians discussed smoking with only half of their smoking patients and infrequently made follow-up appointments about smoking cessation.[5,6] To determine whether greater involvement of administrative and office staff, and provision of specified follow-up time, would increase physician counseling, we conducted a controlled trial.

From the Division of General Internal Medicine, Department of Medicine, University of California, San Francisco (Duncan, Cummings); Kaiser Permanente Medical Center, San Francisco (Stein); and the Department of Epidemiology and International Health, UCSF (Cummings). Address reprint requests to Carol Duncan, RD, MPH, Substance Abuse Research Group, San Francisco Veterans Administration Medical Center, Bldg. 203, Room 3B-4, 4150 Clement Street, San Francisco, CA 94121-SFVAMC 116W. This paper, submitted to the Journal January 16, 1990, was revised and accepted for publication October 24, 1990.

TABLE 1—Baseline Characteristics of Patients in Sample Groups

Characteristics	Training-Plus Before	Training-Plus After	Training-Only
Number of patients:	102	105	139
Female (%)	28	36	43
Mean age (years)	45	46	48
College graduates (%)	34	34	27
Mean cigarettes/day	18	16	17
Occasional smoker (%)	11	18	9
Want to quit smoking (mean, 10-pt. scale)	7.0	7.2	7.1
Confidence in quitting (mean, 10-pt scale)	5.0	5.9	5.2
First visit to this doctor (%)	4	28	27
Type of visit (%)			
Routine follow-up	44	53	50
Urgent problem	33	19	17
Check-up/physical	21	27	29
Other	2	1	4

Methods

We studied two physically separate medical stations at a Kaiser-Permanente Medical Center in the San Francisco Bay Area where internists had not been previously trained in smoking cessation and each saw a minimum of 20 patients per week. Four internists at one medical station agreed to receive training with nursing and administrative support (training-plus), while five internists at another station agreed to receive the training (training-only). One training-plus physician subspecialized in pulmonology; one training-only physician subspecialized in endocrinology, and another in infectious disease. One physician in the training-plus group currently smoked. Two physicians in each group were female, and the mean ages of the two groups were similar (36 years for training-plus; 34 years for training-only).

Research staff met with administrators, then with nurses and their supervisors, then with training-plus physicians and nurses to plan the system of administrative supports. Nurses agreed to ask patients if they smoked, to mark the charts of smokers, and to ask smokers to fill out smoking history forms. Physicians decided to use pre-stamped Nicorette prescription pads. The Department of Medicine designated two appointment slots per week per provider for follow-up of smokers. The group adopted a follow-up protocol using a two-part written Quit-date prescription: one copy was given to the patient and one copy was filed by quit date for nurses to use to telephone patients on quit dates.

All providers concurrently attended two one-hour seminars designed to demonstrate and practice counseling about smoking cessation. These seminars have been described in detail.[5] Training for the training-plus group differed in that nurses and their supervisors attended the sessions. Physicians in the control station were encouraged to ask their office staffs to use reminder stickers to identify smokers.

TABLE 2—Comparison of Physician Counseling About Smoking Between Trained Groups

–	Training-only	Training-plus	Difference	(95% C.I.)
(%)				
Smoking discussed	48.9	68.7	19.8	(6.5,33.2)
Physician initiated discussion of smoking	27.1	62.6	35.5	(22.5,48.6)
Physician asked for quit date	15.8	28.3	12.5	(0.8,24.2)
Smoker agreed to date	7.5	18.2	10.7	(1.0,20.4)
Smoker received written reminder of date	2.3	15.2	12.9	(4.5,21.3)
Nicotine gum prescribed	6.8	15.3	8.5	(−0.7,17.7)
Referred to program	10.5	25.5	15.0	(4.0,26.0)
Received self-help booklet	9.8	34.3	24.6	(13.1,36.1)
Given follow-up appointment about smoking	9.0	18.8	9.8	(−0.4,19.8)

*Smallest sample sizes: 131 for training-only and 96 for training-plus due to missing data.

We defined a smoker as anyone who had smoked a tobacco cigarette during the past seven days. Research staff surveyed patients in the waiting areas, and asked all smokers to participate. This continued until a minimum of 15 smokers were enrolled per physician. Participants answered a questionnaire about their smoking habit, desire to quit and confidence in ability to quit. As soon as possible after the visit, an interviewer telephoned each smoker to determine if the physician or staff had discussed smoking with them and, if so, what steps the physician had recommended. These data were first collected before any meetings with the training-plus group. The week after training was completed data collection began simultaneously from both training-only and training-plus groups' patients.

To monitor use of designated follow-up appointments for smoking we tallied computerized visit records for "smoking" appointments. We also verified with physicians which of their appointments had actually been used for counseling smokers.

Since outcomes were measured by patient report, patients were the unit of analysis. Statistical significance was tested using t-test for continuous and chi-square for categorical data. We also used logistic regression to adjust for differences in outcomes between the groups, for baseline differences in demographic characteristics in patients, their confidence in quitting, occasional smoking status and type of visit. Because these adjustments did not affect results, we present unadjusted differences in the tables.

Results

We enrolled 335 patients; 26 per physician both before and after training in the training-plus group, and 28 per physician in the training-only group. Telephone interviews were completed for 96 percent of patients within 72 hours of their visits. Patients in the three study groups had similar characteristics, but those enrolled during the pre-training period were less confident of being able to quit, were less often making first visits and more often had urgent problems than those enrolled during the post-training period (Table 1). Compared with smokers in the training-plus group, more patients in the training-only group were regular smokers.

Training-plus physicians discussed smoking with their smoking patients more often than training-only physicians, initiated the discussion more often, asked for more quit dates, gave more written reminders and booklets, referred more patients to groups, programs or professional help, prescribed nicotine gum more often, and made more follow-up appointments than training-only physicians (Table 2). Individual physicians in the training-plus group counseled from 52 to 93 percent of patients, while those in the training-only group counseled from 32 to 63 percent of patients. Office staffs of training-plus physicians discussed smoking with 5.2 percent of patient smokers while office staffs of training-only physicians discussed smoking with none.

Comparing their performance from before the program started to after they received training and office support, physicians in the training-plus group counseled more patients (from 38 to 69 percent), initiated the discussion more often (25 to 63 percent), suggested quit dates more often (3 to 28 percent), prescribed nicotine gum more often (4 to 15 percent), referred more patients to programs (11 to 26 percent), gave more booklets (12 to 34 percent), and made more follow-up appointments (10 to 19 percent). All differences were statistically significant ($P < .05$ for all comparisons).

Forty (60 percent) of 68 available smoking appointments were filled, and 23 (34 percent) of designated appointments were actually used for smokers. Only five (22 percent) of these 23 smokers failed to keep their appointments.

Discussion

In this pilot study we found that involvement of staff and administrators in planning and support of smoking cessation counseling enhanced the effect of physician training. Solberg and Kottke have suggested that it is important to enlist office staff in support of smoking cessation efforts in medical practices[7] and the National Cancer Institute has emphasized involvement of staff and administration in smoking cessation in its most recent booklet for health professionals. We found this approach to be quite acceptable to these volunteers, but it is rarely used in practice. For example, we have found that 60 percent of health maintenance organization physicians never have other staff counsel and 68 percent never make follow-up appointments about quitting.[3]

As expected, the official designation of time for follow-up visits substantially increased follow-up of smokers. Not all appointments were used for smokers, but the total number (23) may have been appropriate since physicians were advised to make appointments with smokers who agreed to quit dates, and 69 smokers had so agreed. There is concern that smokers may frequently fail to keep follow-up appointments about cessation, but the proportion of missed appointments by smokers was within the 15–33 percent range reported for all patients in adult medical practices.[8–10]

This was a pilot study. Our small number of physician subjects were volunteers who were not randomized, and telephone interviews were not done blindly. The end point was physician counseling, not smoking cessation. Larger and longer-term randomized trials are needed to determine whether organizational involvement and follow-up appointments increase rates of smoking cessation among patients as well as to identify which components are most effective.

We conclude that systematic staff support and designated follow-up visits for smokers increase smoking cessation efforts of physicians beyond those that result from training alone. Whether these interventions increase cessation rates deserves further study. ☐

Acknowledgments

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Tobacco Consumption Declining Faster in Canada Than in the US

Murray J. Kaiserman, PhD, and Byron Rogers, BA

ABSTRACT

Consumption and prices of tobacco products in Canada and the United States are compared over the period 1981 to 1989. In Canada, consumption has fallen from a higher level and at a faster rate than in the United States (3.52 kg/adult to 2.48 kg/adult vs 3.21 kg/adult to 2.48 kg/adult). With respect to cost, the real tobacco price index in Canada has grown faster than its US counterpart. These results are attributed to higher taxation practices, vigorous legislation, and the leadership role adopted by the Canadian government during the 1980s. (*Am J Public Health*. 1991:81:902–904)

Can a vigorous national policy that includes high taxation and strong legislation towards smoking bring about significant declines in tobacco consumption? In Canada, the answer is yes.

The most recent figures on total tobacco product consumption expressed as kg of tobacco per person aged 15 years and older show that Canada's consumption fell from a level that was higher than that of the United States in 1981 to the same, lower level by 1989 (Table 1). The choice of the age of 15 years reflects the fact that adult smoking patterns are typically established between the ages of 15 and 18 years.[1,2]

In 1981, Canadians aged 15 or older, consumed 3.52 kg of tobacco per capita.[3] Their American cousins, on the other hand, consumed 3.21 kg per capita in 1981, or 8.8% less.[4]

By the end of 1989, Canadians, like Americans, consumed 2.48 kg annually. For the Canadians, this represented a decrease of 22.8% in total weight of tobacco products consumed. By the end of the decade, the Canadian per capita tobacco consumption rate had declined to reach the same level as its United States counterpart (Figure 1).

Over the 9-year period, the slope of the decline in per capita tobacco consumption in Canada has been about 30% steeper than the trend in the United States. Statistical analysis with a multiregression model shows that the difference in decline between Canada and the United States is statistically significant and is probably attributable to policies within the country rather than external influences.

One of the determinants in short-term tobacco demand is the real price of tobacco products, that is, the price adjusted for inflation.[4,5] Graphs of the real tobacco price index and tobacco consumption for both countries show the expected inverse relationship (Figures 2 and 3).

Canada's steeper tobacco price index curve is indicative of that country's greater willingness to tax tobacco, both federally and provincially (Figure 4). In Canada, federal and provincial taxes together account for approximately two thirds of the retail price of a pack of cigarettes. In the United States, federal and state taxes account for approximately 20% of the retail price.[1,4]

In the 1989 Canadian Federal Budget, the finance minister cited health reasons, in addition to fiscal considerations, in justifying a $0.35 (US) increase per pack of 20 in federal cigarette taxes. Provinces impose their own tobacco taxes in addition to the federal tax, thus further raising the real price of tobacco products. In Canada, the average federal and provincial tax rate is now about $2.25 (US) per pack of 20 whereas in the United States the average federal and state tax rate is $0.38 (US) per pack of 20.[4]*

In addition, the 1989 Canadian Federal Budget reduced the tax rate gap between fine-cut tobacco or roll-your-own (RYO) and manufactured cigarettes. Preliminary sales data reported to Statistics Canada indicate that, in the first half of 1990, consumption of RYO tobacco, long the choice of those Canadian smokers who preferred not to quit in response to cigarette price increases, is now also declining. Cigarette rolling tobacco accounts for 14% of the Canadian cigarette market, but only about 1% of the market in the United States.[1,4]

A factor in Canada's relatively recent trend toward higher tobacco tax rates has been the federal government's adoption of a comprehensive approach to discouraging tobacco use by Canadians. Key elements of the National Strategy to Reduce Tobacco Use include a ban on tobacco product advertising, restrictions on promotional activities, prominent new health warning

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Murray J. Kaiserman and Byron Rogers are with the Tobacco Products Section, Environmental Health Directorate, Health and Welfare Canada.

Requests for reprints should be sent to Murray J. Kaiserman, Tobacco Products Section, Environmental Health Centre, Health and Welfare Canada, Ottawa, Ontario, K1A 0L2.

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*On Feb. 26, 1991 the finance minister increased the federal tax by 52 cents per pack of 20. After subsequent provincial budgets, the average tax rate is now about $3.24 per pack of 20 or about 75% of the retail price.

TABLE 1—Real Tobacco Price Indexes and Per Capita Consumption of Tobacco Products Consumed in Canada and the United States Between 1981 and 1989

Year	Real Tobacco Price Index in Canada[5]	Per Capita Consumption in Canada	Real Tobacco Price Index in United States[4]	Per Capita Consumption in United States[4]
1981	100.0	3.52	100.0	3.21
1982	104.1	3.51	104.8	3.17
1983	115.7	3.32	121.4	3.06
1984	123.6	3.19	124.2	2.92
1985	139.5	3.08	127.0	2.91
1986	157.4	2.90	133.1	2.80
1987	163.7	2.78	137.6	2.68
1988	173.0	2.69	144.2	2.62
1989	199.3	2.48	155.1	2.48

messages on tobacco product packaging, health promotion and community support activities around smoking prevention and cessation, economic adjustment programs to assist tobacco growers leaving the industry, and bans or restrictions on smoking in workplaces and on common carriers within federal jurisdiction.

Internationally, Canada has taken a leadership role in banning tobacco advertising, increasing the prominence and legibility of health warning labels, and restricting smoking on longer-duration commercial airline flights.[6] Like Canada, the United States has now banned smoking on flights of up to 6 hours duration, and

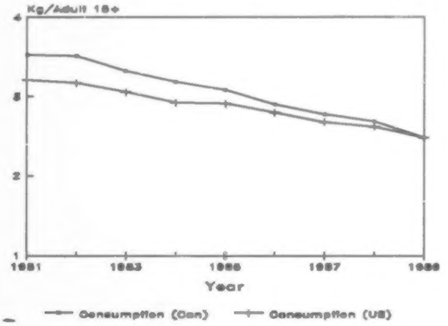

FIGURE 1—Comparison of per capita tobacco consumption in the United States and Canada (1981–1989)

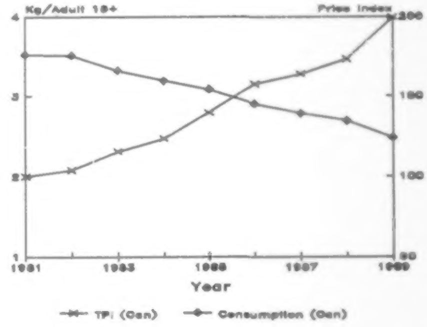

FIGURE 2—Tobacco consumption and tobacco prices in Canada (1981–1989)

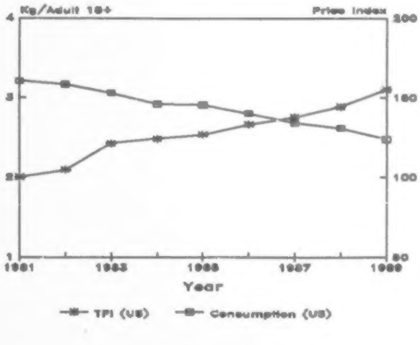

FIGURE 3—Tobacco consumption and tobacco prices in the United States (1981–1989)

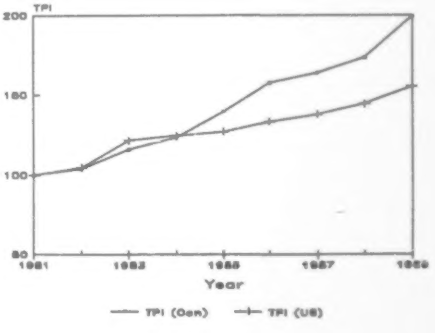

FIGURE 4—Comparison of real tobacco price indices in the United States and Canada (1981–1989)

the US Congress is considering proposals to further restrict or ban cigarette advertising.

For policymakers, the conclusions are simple. Although education and smoking cessation programs can result in moderate declines in tobacco consumption, major declines can be achieved through increased taxation and advertising bans. As a result of an across-the-board strategy to reduce tobacco use, Canada, whose citizens were once characterized as "champion smokers," has caught up with and appears poised to overtake the United States in the trend toward butting out. ☐

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The Impact of Time in Treatment on the Employment and Earnings of Drug Abusers

Michael T. French, PhD, Gary A. Zarkin, PhD, Robert L. Hubbard, PhD, and J. Valley Rachal, MA

ABSTRACT

We use data from a longitudinal survey to estimate the effects of time in drug abuse treatment on post-treatment weeks worked and earnings for 2,420 clients in three treatment modalities. The regression analysis shows that time in treatment had a positive and statistically significant impact on these labor market outcomes, but the effects were small for all modalities. Although residential clients experienced the largest relative changes in weeks worked and real earnings, a benefit-cost calculation suggests that additional residential treatment cannot be justified from earnings improvements alone. These results may indicate a need for more employment services while in treatment. (*Am J Public Health.* 1991; 81:904–907)

Introduction

Although recent studies have shown that drug abuse treatment is successful in lowering drug use and criminal activity for some clients,[1–6] little is known about the impact of treatment on clients' employment and earnings.[7] This paper uses multivariate regression models to examine the impact of time in treatment on post-treatment weeks worked and real (inflation-adjusted) earnings of over 2,400 drug abusers.

Methods

We use data from the Treatment Outcome Prospective Study (TOPS), a long-term, large-scale study of 11,000 drug abusers who entered treatment in 1979, 1980, and 1981 at 41 selected drug abuse treatment programs.[4] It provides extensive information about the nature of drug abuse treatment and describes characteristics and behavior of drug abusers prior to treatment, at admission, and at several follow-up intervals. The sample for this study includes all 2,420 clients who participated in the 12-month follow-up interview—the largest follow-up cohort in TOPS. Hubbard and colleagues[4] report and compare sociodemographics at intake and follow-up from a TOPS program for clients in each modality. Rather than repeating those descriptive statistics here, we refer the reader to their book.

We develop regression models to understand the variation in employment and earnings outcomes. For example, consider a simple regression model that describes a labor market outcome as a function of several explanatory variables[8,9]:

$$Y_f = X_f\beta + \gamma PRETOPS + \delta TOPS + \phi POSTTOPS + \varepsilon, \qquad (1)$$

Address reprint requests to Michael T. French, PhD, Senior Economist and Section Head, Health and Human Resource Economics, Research Triangle Institute, P.O. Box 12194, 3040 Cornwallis Road, Research Triangle Park, NC 27709-2194. The co-authors are also at Research Triangle Institute, where Dr. Zarkin is Senior Economist, Health and Human Resource Economics; Dr. Hubbard is Senior Psychologist and Program Director, Alcohol and Drug Research; Rachal is Director, Center for Social Research and Policy Analysis. This paper, submitted to the Journal July 12, 1990, was revised and accepted for publication January 8, 1991.

*Because methadone maintenance usually involves long-term treatment, some individuals were still in a TOPS treatment program at the time of their follow-up interview. For these clients, the period "one year following discharge" actually refers to the period starting one year following admission.

TABLE 1—Regression Coefficient Estimates for the "Real Earnings" Specification by Treatment Modality[a]

	Modality		
Variables	Methadone Maintenance (N = 737)	Residential (N = 655)	Outpatient Drug-Free (N = 750)
Intercept	$-14591.00^{\dagger\dagger}$	-2445.50	$-11573.00^{\dagger\dagger}$
	(2460.93)	(1637.52)	(2217.65)
Males	$2918.51^{\dagger\dagger}$	$1679.90^{\dagger\dagger}$	$4113.92^{\dagger\dagger}$
	(633.08)	(534.57)	(651.68)
Caucasian/white	$3351.42^{\dagger\dagger}$	$2034.48^{\dagger\dagger}$	$1953.35^{\dagger\dagger}$
	(679.22)	(516.64)	(780.67)
Highest grade attended	$965.71^{\dagger\dagger}$	$247.74^{\dagger\dagger}$	$1133.72^{\dagger\dagger}$
	(129.48)	(102.59)	(138.97)
Age (months)	2.78	3.36	0.78
	(3.52)	(2.69)	(3.76)
Married	$1209.12^{\dagger\dagger}$	736.65	602.12
	(590.49)	(537.15)	(677.57)
Head of household	$1463.22^{\dagger\dagger}$	-77.72	$2112.84^{\dagger\dagger}$
	(604.54)	(456.72)	(645.49)
Number of dependents	$707.14^{\dagger\dagger}$	137.17	$970.03^{\dagger\dagger}$
	(216.64)	(203.18)	(267.93)
Total weeks in treatment prior to entering a TOPS treatment program	-1.47	-2.20	$-14.21^{\dagger}$
	(3.62)	(4.29)	(7.83)
Total weeks in a TOPS treatment program	$21.09^{\dagger\dagger}$	$43.90^{\dagger\dagger}$	14.31
	(9.54)	(8.27)	(14.58)
Total weeks in treatment 12 months after leaving a TOPS treatment program	$-29.80^{\dagger\dagger}$	-18.35	-31.12
	(13.46)	(14.67)	(23.45)
Admitted to a TOPS program in:			
Chicago	912.02	-1190.83	—
	(1125.91)	(893.80)	
New Orleans	$4626.42^{\dagger\dagger}$	-694.88	-1433.59
	(1250.65)	(941.13)	(1465.23)
New York	1316.49	-341.41	$-2214.14^{\dagger}$
	(1007.28)	(1171.37)	(1292.49)
Phoenix	$2263.77^{\dagger\dagger}$	$-2025.50^{\dagger\dagger}$	-1146.80
	(1150.48)	(1016.83)	(1107.85)
Portland	—	-1155.11	-1186.32
		(965.59)	(1167.57)
Miami	1708.15	550.36	58.88
	(1313.09)	(948.43)	(1298.62)
San Francisco	--	-385.33	$-5519.56^{\dagger\dagger}$
		(982.57)	(1553.10)
R-square	0.21	0.13	0.22

[a] Dependent variable is total real eamings (1983 dollars) from full- or part-time jobs during the 12 months after leaving a TOPS program.

†Significantly different from zero at the 90% confidence level (two-tailed test).

††Significantly different from zero at the 95% confidence level (two-tailed test).

Note: Standard error in parentheses.

where Y_f is a labor market outcome variable at the one-year follow-up, X_f is a vector of explanatory variables at the one-year follow-up, *PRETOPS* is a measure of total time in treatment prior to entering a TOPS program, *TOPS* is total time in a TOPS program, and *POSTTOPS* is total time in treatment during the year after leaving a TOPS program.* From Equation (1) we can estimate the effect of TOPS treatment on a labor market outcome variable, controlling for demographics and prior and ensuing treatment episodes.

Equation (1) may yield biased parameter estimates for the treatment duration variables if unobserved client differences (e.g. motivation, discipline, aggression) are correlated with the time in TOPS treatment.[10] To demonstrate this situation conceptually, consider Equations (2) and (3), which explain labor market variable Y at admission to a TOPS program (Y_a) and one year after leaving a TOPS program (Y_f), respectively:

$$Y_a = X_a\beta + \gamma PRETOPS + v + \varepsilon_a \qquad (2)$$

$$Y_f = X_f\beta + \gamma PRETOPS + \delta TOPS + \phi POSTTOPS + v + \varepsilon_f \qquad (3)$$

where v is an error term. If v includes unobserved client characteristics, single-equation least-squares regressions may yield biased coefficient estimates. To correct this potential problem, we propose a "first-difference" model as illustrated in Equation (4):

$$(Y_f - Y_a) = (X_f - X_a)\beta + \delta TOPS + \phi POSTTOPS + (\varepsilon_f - \varepsilon_a). \qquad (4)$$

By differencing the variables at follow-up with the same variables at admission, we eliminate the error component, v, and all other variables that do not change between admission and follow-up. The treatment duration coefficient, δ, now represents an unbiased estimate of the effect of time in a TOPS treatment program on labor market changes.

Results

Table 1 reports regression estimates for the real earnings specification of Equation (1). Coefficient estimates for the time in TOPS treatment variable were positive for each modality and statistically significant for methadone maintenance and residential clients. On average, one additional week in treatment increased annual real earnings for methadone clients by $21 and for residential clients by $44 during the year following TOPS treatment.

The dependent variable in the Table 2 equations is total weeks worked during the 12 months after leaving a TOPS program. Coefficient estimates for time in a TOPS treatment program were positive and statistically significant for every modality. The coefficients for the residential and outpatient drug free equations imply that six to seven additional weeks in treatment led to one additional week of employment. Methadone clients had a much lower estimated return, with 16 additional weeks in treatment needed for each additional week of employment.

In addition to demographic and treatment duration variables, Tables 1 and 2 report coefficient estimates for program location. Differences in city location coefficients reflect geographical differences in labor markets and program-specific differences in treatment delivery.

Table 3 lists coefficient estimates for the first-difference specifications. Several variables "difference out" because they did not change from admission to follow-up (e.g. race, sex, weeks in treatment prior to TOPS). Focusing on the time in treatment variables, we see that total

TABLE 2—Regression Coefficient Estimates for the "Weeks Worked" Specification by Treatment Modality[a]

Variables	Modality Methadone Maintenance (N = 753)	Residential (N = 667)	Outpatient Drug-Free (N = 766)
Intercept	−24.273[††]	10.645[††]	4.635
	(6.066)	(5.394)	(4.994)
Males	7.012[††]	5.243[††]	7.119[††]
	(1.563)	(1.755)	(1.470)
Caucasian/white	6.522[††]	5.663[††]	6.156[††]
	(1.681)	(1.698)	(1.749)
Highest grade attended	2.165[††]	0.815[††]	1.859[††]
	(0.321)	(0.337)	(0.315)
Age (months)	−0.002	−0.002	−0.021[††]
	(0.009)	(0.009)	(0.008)
Married	1.390	1.235	1.433
	(1.459)	(1.767)	(1.538)
Head of household	3.733[††]	−2.268	3.046[††]
	(1.490)	(1.501)	(1.461)
Number of dependents	1.365[††]	0.792	1.389[††]
	(0.532)	(0.674)	(0.609)
Total weeks in treatment prior to entering a TOPS treatment program	−0.015	0.003	−0.043[††]
	(0.009)	(0.014)	(0.018)
Total weeks in a TOPS treatment program	0.061[††]	0.168[††]	0.137[††]
	(0.023)	(0.027)	(0.033)
Total weeks in treatment 12 months after leaving a TOPS treatment program	−0.054	−0.032	−0.032
	(0.033)	(0.049)	(0.053)
Admitted to a TOPS program in:			
Chicago	0.155	−13.963[††]	—
	(2.781)	(2.957)	
New Orleans	7.231[††]	−5.280[†]	−4.[illegible]39
	(3.077)	(3.107)	(3.304)
New York	0.564	−9.432[††]	−10.379[††]
	(2.479)	(3.874)	(2.920)
Phoenix	8.899[††]	−10.995[††]	−4.106
	(2.846)	(3.348)	(2.522)
Portland	—	−11.661[††]	−6.624[††]
		(3.224)	(2.639)
Miami	6.008[†]	−4.696	−0.344
	(3.213)	(3.131)	(2.947)
San Francisco	—	−10.829[††]	−20.119[††]
		(3.260)	(3.539)
R-squared	0.184	0.162	0.213

[a] Dependent variable is total weeks worked at full- or part-time jobs during the 12 months after leaving a TOPS program.

[†] Significantly different from zero at the 90% confidence level (two-tailed test).

[††] Significantly different from zero at the 95% confidence level (two-tailed test).

Note: Standard error in parentheses.

weeks in a TOPS treatment program had a positive impact on real earnings and weeks worked in all equations. Only the coefficient estimate from the real earnings specification for methadone clients was not significantly different from zero.

Discussion

This is the first longitudinal study to analyze the impact of length of stay in drug abuse treatment on labor market outcomes. Regression analyses indicate that time in treatment had a statistically significant positive, but small, impact on the follow-up employment and earnings of clients in the major treatment modalities. We believe this result can be explained, in part, by the dearth of training and employment services that were available to treatment clients during this period. Treatment resources were often directed at addiction control and behavior modification at the expense of vocational services. To reverse this trend, greater investments in vocational and employment services may be necessary to produce a strong and lasting effect on labor market outcomes.

Interestingly, time in treatment prior to and following TOPS treatment had a generally negative, but insignificant, effect on real earnings and weeks worked for each modality. This result implies that individuals with more treatment prior to and following TOPS (holding constant time in TOPS treatment) are less likely to have labor market success. Perhaps treatment episodes before and after TOPS signal chronic treatment and employment problems for these clients.

Overall, residential clients experienced the largest impact from time in treatment. From the regression results, we found that residential clients would have earned approximately $44 more if they stayed an additional week in treatment. But the average weekly cost of residential treatment for each TOPS client was approximately $143.[4] Looking only at the earnings outcome of time in treatment, the investment in additional residential treatment would not be cost-beneficial. However, after factoring in the dollar benefits of other treatment outcomes (e.g. lower drug use, reduced criminal activity, improved health status), residential treatment is likely to generate a higher return on the investment. Further analyses of the treatment process, treatment outcomes, and linkages between criminal activity, drug use, and employment are necessary to help design contemporary drug abuse treatment strategies. ☐

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TABLE 3—Regression Coefficient Estimates for the "First-Difference" Specifications by Treatment Modality

	Dependent Variable – Difference in Real Earnings[a]			Dependent Variable – Difference in Weeks Worked[b]		
Variables	Methadone Maintenance (N = 707)	Residential (N = 626)	Outpatient Drug-Free (N = 714)	Methadone Maintenance (N = 732)	Residential (N = 646)	Outpatient Drug-Free (N = 766)
Intercept	−1477.69	−1634.75††	−200.37	−4.90††	−2.31	−1.29
	(1028.17)	(693.39)	(585.78)	(2.25)	(1.79)	(1.56)
Difference in highest grade attended	731.19††	−3.29	131.28	0.79	0.33	−0.47
	(283.45)	(179.77)	(204.17)	(0.63)	(0.46)	(0.53)
Difference in age	−1.71	26.92	21.04	0.01	0.08	0.06
	(30.77)	(20.74)	(16.39)	(0.07)	(0.05)	(0.04)
Difference in marital status	−247.21	209.12	1851.95††	0.29	0.53	6.13††
	(696.24)	(724.34)	(674.66)	(1.54)	(1.85)	[illegible]
Difference in household status	1036.43†	−1057.83††	675.97††	4.75††	−3.91††	2.93††
	(643.02)	(537.78)	(511.40)	(1.42)	(1.38)	(1.33)
Difference in number of dependents	435.32	275.70	490.32†	2.19††	0.62	1.09
	(276.75)	(277.17)	(272.80)	(0.61)	(0.72)	(0.74)
Weeks in a TOPS treatment program	10.43	34.60††	24.45†	0.06††	0.14††	0.07††
	(12.98)	(12.68)	(13.56)	(0.03)	(0.03)	(0.04)
Weeks in treatment year after TOPS	3.86	−8.11	−23.61	−0.04	0.02	−0.04
	(17.69)	(22.26)	(21.12)	(0.04)	(0.06)	(0.06)
R-square	0.02	0.03	0.03	0.05	0.05	0.04

[a] Dependent variable is total real earnings during the year after leaving a TOPS program minus total real earnings during the year prior to entering a TOPS program.
[b] Dependent variable is total weeks worked during the year after leaving a TOPS program minus total weeks worked during the year prior to entering a TOPS program.
†Significantly different from zero at the 90% confidence level (two-tailed test).
††Significantly different from zero at the 95% confidence level (two-tailed test).
Note: Standard error in parentheses.

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Television Viewing and Obesity in Adult Females

Larry A. Tucker, PhD, and Marilyn Bagwell, RN, PhD

ABSTRACT

We measured the relation between time spent watching television per week and obesity in 4,771 adult females. After controlling for age, education, cigarette smoking, length of work week, and weekly duration of exercise, females who reported three to four hours of TV viewing per day showed almost twice the prevalence of obesity (body fat > 30 percent), and those who reported more than four hours of TV watching per day showed more than double the prevalence of obesity, compared to the reference group (<1 hr/day). Part of the TV/obesity association was a function of differences in exercise duration among the four TV viewing categories. (*Am J Public Health.* 1991;81:908–911)

Introduction

Lifestyle plays a key role in the development and prevention of a host of chronic diseases. Until recently, few have recognized television viewing as a significant lifestyle factor. Yet, next to sleep and work, watching TV consumes more time than any other activity in America.[1,2] Adolescents spend more time watching TV each year than they spend in school, and the typical adult averages nearly four hours of TV viewing per day.[1–3]

Research has determined that frequent television viewing may function as a risk factor in a number of disorders.[4–10] Health problems may result from many hours of daily TV watching due to the deceptive and inaccurate health-related messages which are regularly transmitted to viewers.[10,11] However, television's principal problem is probably more related to the physical passivity which accompanies TV viewing rather than the misleading micro-lessons it teaches.

TV has profoundly changed the use of leisure time in America.[2,12] While people watch TV, physical activity tends to be minimal and snacking is prevalent.[10,13,14] Conditions for hypokinetic ailments such as obesity and poor fitness dominate.

In 1986, Tucker[6] studied the connection between television viewing and multiple measures of physical fitness in adolescent males. Findings showed that high levels of TV watching were strongly associated with low scores on tests of fitness, (pullups, situps, sidestep, six-minute run, and pushups). Similarly, Dietz and Gortmaker[5] found that as TV viewing increased, obesity increased systematically in several thousand children aged 6–11 and 12–17. In 1989, Tucker and Friedman[8] showed that time spent watching TV was directly related to obesity levels in adult males. Men who watched more than three hours of TV per day had more than twice the prevalence of obesity compared to those who watched less than one hour per day. To date, the association between TV viewing and obesity in adult females has not been studied; hence, the present investigation was conducted. A secondary purpose was to ascertain the extent to which age, cigarette smoking, hours worked per week, and weekly exercise duration mediate the relation between TV viewing and obesity in women.

Methods

Subjects

Subjects were 4,771 females employed by over 30 different companies in the United States. All subjects worked for companies which paid to have their employees participate in a health promotion/disease prevention screening program offered by Health Advancement Services, Inc. Participants were distributed nearly equally across family income categories ranging from $10,000 to $60,000+. Median age of the sample was 35 years and approximately 83 percent of the females were White, 60 percent were married, 18 percent separated or divorced, and 50 percent reported some college education.

Instrumentation and Procedures

All data were gathered by registered nurses employed by Health Advancement Services as part of their ongoing screening program. Subjects completed an informed consent form and each was tested individually during a private 50-minute examination. A structured, written questionnaire was used to collect the demographic and lifestyle information, including TV viewing, exercise, and smoking habits. A Harpenden skinfold caliper was employed to assess subcutaneous fat at three body sites on each subject: ***thigh***, vertical fold taken on the anterior aspect midway between the hip and the knee; ***triceps***, vertical fold taken on the back of the arm midway between the shoulder and the elbow; ***iliac crest***, diagonal fold taken 2.5 cm above the iliac crest. The sum of the skinfold measurements along with age were used to estimate the total body fat percentage of each subject.[15] Research shows that this three-site skinfold method of

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Address reprint requests to Larry A. Tucker, PhD, Director of Health Promotion, 273 SFH, Brigham Young University, Provo, UT 84602. Dr. Bagwell is with Arizona State University, Tempe. This paper, submitted to the Journal June 11, 1990, was revised and accepted for publication January 23, 1991.

TABLE 1—Hours of Television Viewing, Body Fat, and Obesity According to the Control Variables

Control Variables	N	Hours of TV per Day (row %) <1	1–2	3–4	>4	Body Fat Percentage (row %) <18	18–23	24–29	30+*
Total Group	4,771	12.3	59.2	24.1	4.4	10.7	26.2	30.9	32.3
Age (years)									
19–29	1,541	10.4	56.5	26.8	6.3	20.9	36.5	25.8	16.8
30–39	1,350	13.5	62.3	21.2	3.0	8.9	29.4	32.7	29.0
40–49	997	17.0	60.1	20.3	2.6	3.1	18.6	35.5	42.8
50–59	654	9.3	62.5	25.5	2.7	2.4	12.4	30.9	54.3
60+	229	5.3	50.2	34.2	10.2	2.3	7.7	33.5	56.6
Highest Education									
elementary	71	11.6	37.7	37.7	13.0	6.1	21.2	25.7	47.0
high school	1,941	8.5	57.3	28.6	5.6	11.6	25.2	30.7	32.6
vocational	357	10.7	55.2	29.6	4.5	7.4	24.3	33.1	35.1
college	1,726	14.6	61.1	20.5	3.9	12.3	29.5	29.9	28.3
professional	676	18.3	63.8	16.4	1.5	6.5	22.5	33.4	37.6
Smoking Status									
never smoked	2,451	13.1	60.1	23.3	3.5	9.9	25.2	29.4	35.6
ex-smoker	881	13.5	60.2	22.5	3.9	8.3	24.5	32.5	34.7
1–20 daily	1,020	9.9	56.2	27.7	6.2	15.3	29.4	32.2	23.1
21+ daily	419	9.3	58.9	27.9	4.0	10.5	31.0	31.5	26.9
Weekly Exercise									
<30 minutes	2,735	10.5	57.5	26.7	5.2	9.9	24.4	29.8	35.9
31–60 min	421	13.2	61.4	22.0	3.4	11.4	27.5	31.7	29.4
61–100 min	426	15.0	59.7	22.0	3.4	7.6	28.4	35.9	28.1
101–200 min	655	14.2	61.0	21.3	3.5	11.4	29.3	33.4	25.9
201+ min	534	16.4	63.0	17.7	2.9	16.2	29.8	29.2	24.8
Work Week									
<20 hrs/wk	215	8.0	43.4	32.6	16.0	8.1	24.4	36.4	31.1
21–40 hrs/wk	2,953	11.0	58.6	26.0	4.4	11.7	27.4	29.7	31.2
41–50 hrs/wk	1,422	14.7	63.0	19.9	2.4	9.7	24.6	32.0	33.7
51+ hrs/wk	181	19.6	60.3	18.4	1.7	6.2	26.4	30.9	36.5

*Females with 30 percent or more body fat were labeled obese.

body fat assessment correlates well with hydrostatic weighing (r = .84).[15] Obesity was defined as 30 percent or more body fat.[16] Body fat, television viewing, and the control variables were categorized as depicted in Table 1. Indirect validation of the body fat,[8,17–20] TV,[4,6–8,19] and exercise duration[17–20] variables has been demonstrated previously.

The relation between duration of TV viewing and body fat was indexed by the Mantel-Haenszel chi-square statistic which shows the extent of the linear association between ordinal-level, categorical variables.[21,22] The relation between TV viewing and obesity was measured by the odds ratio[23] with subjects who reported less than one hour of TV watching per day used as the reference group. Cochran-Mantel-Haenszel summary risk estimates were used to assess the TV/obesity relation with the potential confounders controlled.[24]

Results

Almost 60 percent of the subjects reported watching one to two hours of TV per day, while approximately one in four indicated three–four hours of daily TV viewing (see Table 1). More than four hours of TV viewing per day was reported most often by subjects who were young or old, by the uneducated, by smokers, by non-exercisers, and by those with short work weeks. Obesity, measured in nearly one-third of the adult females, was more common among older subjects, the uneducated, never and ex-smokers, non-exercisers, and those with long work weeks.

Table 2 shows the estimated risk of obesity by TV viewing time without adjustment and with adjustment for the control variables displayed in Table 1. After adjustment for all of the potentially confounding variables, subjects who reported four+ hours of daily TV watching showed more than two times the prevalence of obesity, and subjects who reported three–four hours of daily viewing showed almost twice the prevalence of obesity compared to females who watched less than one hour of TV daily (see Table 2).

Age had the strongest confounding effect on the TV/obesity relation of any single control variable. After adjustment for age only, estimated risk of obesity was increased by an average of 15 percent for the three TV viewing groups compared to the reference group. Control of exercise duration, in addition to the other potentially confounding factors, also affected the TV/obesity association. Specifically, estimated risk of obesity was decreased by approximately 18 percent for the four+ hour/day viewers and 12 percent for the three–four hour/day viewers compared to the reference group.

Discussion

According to the results, adult females who watch four or more hours of TV per day tend to have more than double the prevalence of obesity compared to those who watch less than one hour daily, as in adult males[8] and children.[5] However, because of the correlational nature of this study, cause-and-effect conclusions are not warranted. Although excessive TV viewing may actually contribute to obesity in women, it is plausible that obesity leads to an abundance of TV watching. As with adult males,[8] females who are fat may be

TABLE 2—Estimated Risk of Obesity by Television Viewing Group

Daily TV Viewing	Variable Controlled	Obese* N	Obese* %	Obese* RR_{cmh}	Obese* 95% CI
<1 hour (n = 585)	none	140	23.9	1.00*	—
1–2 hours (n = 2,825)	none	896	31.7	1.48	1.20, 1.81
	age, education, smoking, work week			1.46	1.15, 1.86
	above + exercise duration			1.29	0.99, 1.68
3–4 hours (n = 1,151)	none	402	34.9	1.71	1.36, 2.14
	age, education, smoking, work week			1.98	1.51, 2.61
	above + exercise duration			1.89	1.37, 2.61
>4 hours (n = 210)	none	73	34.8	1.69	1.20, 2.38
	age, education, smoking, work week			2.62	1.51, 4.53
	above + exercise duration			2.15	1.15, 4.01

Note: RR_{cmh} = Cochran-Mantel-Haenszel summary estimate of relative risk.
*Subjects who were found to have 30 percent or more body fat were labeled obese.

more inclined to watch TV as a primary source of recreation because of the minimal exertion required, whereas non-obese women may prefer other pastimes which demand greater physical involvement.

The strong association between TV watching and obesity in females could be the function of a "third variable" or confounding factor. Each of the control variables of the present study was strongly related to both TV viewing and obesity, except smoking, which was strongly associated with obesity only. Simultaneous adjustment for differences in age, education, length of work week, smoking, and weekly exercise duration substantially increased the estimated risk of obesity for the television watchers, especially the high-duration viewers.

Duration of weekly exercise had a meaningful impact on the TV/obesity relation. High-duration TV viewing was 80 percent more common and obesity was 45 percent more common among non-exercisers compared to those in the highest exercise group (see Table 1). Apparently, part of the reason women who watch a lot of TV tend to be more obese is because they tend to exercise less. If these women spent more time exercising, perhaps in place of some TV viewing time, risk of obesity would likely be decreased.

Other factors not measured in this investigation could confound the TV/obesity association. Diet is a good example. Because television watchers digest thousands of commercial and program messages which encourage consumption of non-nutritious foods,[25–30] and because snacking and eating foods which are advertised on TV are closely linked with television viewing,[10,13,14] it is possible that high-duration TV watchers are more obese than low-duration TV viewers because they consume more food, particularly more junk food, than their counterparts.

Perhaps the most cogent explanation of the association between TV viewing and obesity is based on a dynamic model: as TV viewing time increases, exercise tends to decrease and snacking tends to increase. As exercise decreases and snacking increases, obesity tends to increase. And as obesity increases, attraction to passive recreation, such as watching TV, tends to increase—a cycle of mutual interaction and reinforcement. Future research relative to the TV/obesity relation will need to consider the dietary intake of subjects as well as their physical activity levels. Furthermore, because subjects in this sample were mostly White and well-educated, generalization to other populations will require additional study.

Without question, watching television is a powerful and pervasive lifestyle factor in our society. Given the enormous amounts of time devoted to this idle pastime and the many subtle and distorted health-related messages conveyed by the medium, it is not surprising that children and adults are influenced by TV to a degree far surpassing earlier beliefs.[1–3] Greater effort needs to be directed toward the study of television's influence on health and its role as a precursor of disease. ☐

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Characteristics of a Random Sample of Emergency Food Program Users in New York: I. Food Pantries

Katherine L. Clancy, PhD, Jean Bowering, PhD, and Janet Poppendieck, PhD

ABSTRACT

Food pantry users throughout New York State were studied and many demographic differences found between New York City and Upstate New York respondents. Seven percent of households had no income and median income as percent of the poverty level was 59 percent. Slightly more than 40 percent were spending over 60 percent of their incomes on housing. The data from this survey, the first in New York State to employ a random sampling design, demonstrate a sizable gap between household needs and available resources. (*Am J Public Health*. 1991;81:911–914)

Introduction

Food pantries and soup kitchens have become symbols of hunger and poverty in the United States. Food drives are an accepted feature of life and the Emergency Feeding System (EFS), once regarded as a response to a temporary situation, has become institutionalized. As this occurs, government agencies, scholars, hunger activists, and emergency food providers have begun to question the advisability of expanding or even maintaining the system. Many issues are involved in the emerging debate, including the needs of those using the system, which was the focus of our studies.

New York is one of a few states that provides funding for the operation of its EFS which is composed of food pantries, soup kitchens, and food banks. In 1987 there were 414 pantries in New York City (NYC) serving 2.9 million people and 1,143 in upstate New York plus Long Island (Upstate) serving 2.4 million.[1] This and the following report[2] present data from a survey of persons utilizing the food pantry and soup kitchen components of the EFS throughout the state.

Methods

The food pantry survey was conducted from October 1988 to February 1989, at a random sample of sites drawn from the Bureau of Nutrition census.[1] At each site, selection of individual clients was as random as conditions permitted. The final sample included 311 clients from 29 pantries Upstate and 208 clients from 19 sites in NYC.

The interviewers, mainly students in social science and nutrition, were trained at Syracuse University or Hunter College. The 20-minute interview included primarily closed-ended questions about client demographics, sources and amounts of income, and participation in food and other assistance programs.

The census had shown that the numbers of people served per pantry ranged from a few to 128,000 per year. To verify that our random sampling procedure had preserved this diversity, we found no significant difference between the percentages of clients coming from larger or smaller sites in either Upstate ($\chi^2 = 0.547$, df = 1, p > 0.3), or NYC ($\chi^2 = 0.040$, df = 1, p > 0.8).

From the Department of Nutrition and Food Management, Syracuse University, Syracuse, NY (Clancy and Bowering); and the Department of Sociology, Hunter College, New York, NY (Poppendieck). Address reprint requests to Katherine L. Clancy, PhD, Department of Nutrition and Food Management, 034 Slocum Hall, Syracuse University, Syracuse, NY 13244. This paper, submitted to the Journal May 30, 1990, was revised and accepted for publication January 22, 1991.

TABLE 1—Characteristics of Food Pantry Clients

	% of Responses			
	Upstate (n = 311)	New York City (n = 208)	χ^2 Value	p
Female respondent	67	61	1.847[a]	0.174
Age of respondent (years)				
35 or under	50	34		
36–59	37	37		
60 or older	13	29	24.100[b]	0.0001
Ethnicity				
White	57	16	85.575[a]	0.0001
African-American	31	65	60.702[a]	0.0001
Hispanic	11	17	3.911[a]	0.048
Living/employment situation				
Live alone	23	39	14.950[a]	0.0001
Household with children	61	35	32.077[a]	0.0001
Someone in household employed	29	22	3.535[a]	0.060

[a] χ^2 (df = 1).
[b] χ^2 (df = 2).

TABLE 2—Income Sources of Food Pantry Clients

	% of Households[b]			
	Upstate	New York City	χ^2 Value[a]	p
Wages	24	20	1.268	0.260
Other public assistance	22	13	7.096	0.008
Social Security	21	26	2.096	0.148
SSI/other disability	20	14	2.597	0.107
AFDC (Aid to Families with Dependent Children)	21	16	1.869	0.172
Child support/alimony	8	1	12.582	0.0001
Retirement/pension	6	2	3.909	0.048
Home relief	5	7	0.576	0.448
Veteran's benefits	3	3	0.046	0.831
Unemployment benefits	2	2	0.135	0.713

[a] χ^2 (df = 1).
[b] Percentages do not add up to 100% because more than one response was possible.

TABLE 3—Income Range and Relation to Poverty Level

	% of Responses			
Household Income[a] (Monthly)	Upstate (n = 293)	New York City (n = 188)	Test Statistic	p
\$0	7	8		
\$1–300	22	25		
\$301–600	35	45		
\$601–1000	24	16		
>\$1000	12	6	11.380[b]	0.023
Income < 50% of poverty	41	44	0.424[c]	0.515
Income < 100% of poverty	83	80	0.682[c]	0.409
Median % of poverty	59%	59%	−0.157[d]	0.875
Median monthly per capita income	\$162	\$177	1.255[d]	0.209

[a] Excludes food stamps.
[b] χ^2 (df = 4).
[c] χ^2 (df = 1).
[d] Wilcoxon Rank Sum Test.

Results

As shown in Table 1, more of the NYC sample was age 60 or older. Within the older age group, there was no difference between geographic areas in the percentage of persons living alone; however, among respondents ages 36–59, a larger percentage lived alone in NYC than Upstate.

Income sources are presented in Table 2. Approximately 50 percent of the Upstate sample received public assistance, while 36 percent of the NYC households did so. Household income distributions (Table 3) were significantly different due to a larger percentage of households with higher total incomes Upstate. Approximately 40 percent of each sample had incomes less than 50 percent of poverty.

Sixty-two and 52 percent of Upstate and NYC respondents, respectively, were on Medicaid (χ^2 = 5.030, df = 1, p = 0.025). Approximately one-third of each group received a housing subsidy. Despite this, 45 percent of Upstate and 41 percent of NYC households spent more than 60 percent of their incomes on housing (shelter and utilities).

There was a highly significant difference (χ^2 = 18.286, df = 3, p = 0.0001) in length of time that households had utilized food pantries due to both more long-term users (>3 years) Upstate and more short-term users (<6 months) in NYC. Table 4 shows dissimilarities in households' use of other food programs with regard to use of soup kitchens, school meal programs, and senior meals.

Table 5 presents selected demographic characteristics of food stamp users. In both groups, neither living alone or living with others, nor educational level were related to food stamp use. Upstate, significantly more females, more households with children, and fewer seniors were on the program. In NYC, only employment status was significantly different.

The monthly food stamp benefit per household was significantly different ($129 Upstate and $95 NYC) and the Upstate per capita benefit was also significantly higher. Both samples showed significant differences between food stamp users and non-users with regard to income as a percentage of poverty level (Table 6).

Discussion

Since federal poverty income thresholds are based upon the cost of a mini-

TABLE 4—Use of Other Food Programs by Food Pantry Clients

	% of Households			
	Upstate (n = 311)	New York City (n = 209)	χ^2 Value^{a}	p
All households^{b}				
Food stamps	57	50	1.875	0.171
Soup kitchens	19	40	25.157	0.0001
Free food	41	50	4.079	0.043
Households with people over 60	(n = 60)	(n = 72)		
Meals at senior center	5	32	15.022	0.0001
Home-delivered meals	5	7	0.263	0.608
Households with children under 18	(n = 191)	(n = 78)		
School breakfast (free/reduced)	36	54	5.573	0.018
School lunch (free/reduced)	63	72	1.628	0.202

[a] χ^2 (df = 1).
[b] Percentages do not add up to 100% because more than one response was possible.

mally adequate diet, households below the poverty level are, by definition, at risk of being unable to purchase such a diet.[3] The median income of all participating households was 59% of poverty. Furthermore, federal poverty guidelines were developed when shelter costs were far lower, and the large percentage of income now spent by households further reduces the amount available for food.

Some differences were noted between Upstate and NYC respondents with regard to the extent of use of food programs. People Upstate had utilized pantries for a longer period of time, and there was a positive significant relationship between food stamp participation and length of food pantry use. Food stamp users also made greater use of other food programs. These tendencies suggest that these households are dependent on the EFS even though they are making use of the welfare system in a fairly inclusive manner. In NYC there was not, in most cases, significantly greater use of other programs by food stamp recipients, although the actual percentages of those using soup kitchens, senior meals, etc, was often higher than in the Upstate sample.

Receipt of welfare or food program benefits does not preclude the need for emergency food. Nearly half of respondents were receiving public assistance, but in New York State in 1988 AFDC plus full food stamp benefits brought a household to only 89% of the poverty level.[4] In the Upstate and NYC samples, respectively, 38 percent and 40 percent of respondents 65 and over lived on only Social Security payments which places them near the poverty level.[5] Nearly one-fourth of households included a wage-earner, but a full time minimum wage job brought a household of three to 74 percent of poverty and a household of four to only 57 percent of poverty.[6]

The percentage of households having incomes below 50 percent of the poverty level is similar to the nationwide figure.[7] These households more frequently had children, and, not surprisingly, were less likely to be employed. In both locations, non-White households were significantly more likely to have an income below 50 percent of poverty.

Our findings portray a population in need, and at risk of the health hazards

TABLE 5—Demographic Characteristics of Food Pantry Clients According to Food Stamp Use

	Upstate				New York City			
	+ FS (n = 176)	−FS (n = 135)	Test Statistic	p	+FS (n = 105)	−FS (n = 103)	Test Statistic	p
			% of Responses					
Female respondent	63	37	8.925^{a}	0.003	55	45	2.806^{a}	0.094
Age of respondent (years)								
35 or under	61	39			53	47		
36–59	61	39			51	49		
60 or older	26	74	17.067^{b}	0.0001	44	56	1.107^{b}	0.575
Living/employment situation								
Live alone	61	39	0.779^{a}	0.377	44	56	1.934^{a}	0.164
With children	64	36	12.279^{a}	0.0001	59	41	3.144^{a}	0.076
Someone employed	38	62	17.989^{a}	0.0001	24	76	15.572^{a}	0.0001
Below 50% of poverty	46	34	4.053^{a}	0.044	50	37	3.250^{a}	0.071
Below 100% of poverty	93	70	26.552^{a}	0.0001	90	71	10.651^{a}	0.001
Median % of poverty	51%	74%	3.629^{c}	0.0003	50%	68%	1.947^{c}	0.051
Average monthly food stamp Benefit:								
(Household)	\$129 ± 80^{d} (\$107–149)^{f}	—	—	—	\$95 ± 76^{d} (\$69–121)^{f}	—	−3.467^{e}	0.0006
(Per capita)^{c}	\$43 ± 4^{d} (\$39–47)^{f}	—	—	—	\$37 ± 5^{d} (\$32–42)^{f}	—	2.155^{e}	0.032

[a] χ^2 (df = 1).
[b] χ^2 (df = 2).
[c] Wilcoxon Rank Sum Test.
[d] Standard Deviation.
[e] t-value comparing Upstate vs NYC.
[f] 95% Confidence Interval.

associated with poor nutrition. In concert with similar data from non-random studies done in New York State between 1985 and 1987,[8–10] and additional analyses of this study,[11] a picture emerges of a sizable gap between household needs and available resources, a gap that food pantries have helped to bridge. If private food assistance diminishes (as has been predicted), government aid to the poor must be expanded or the risk of decreased food security and poor nutrition will increase. ☐

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Characteristics of a Random Sample of Emergency Food Program Users in New York: II. Soup Kitchens

Jean Bowering, PhD, Katherine L. Clancy, PhD, and Janet Poppendieck, PhD

ABSTRACT

A random sample of soup kitchen clients in New York City was studied and specific comparisons made on various parameters including homelessness. Compared with the general population of low income persons, soup kitchen users were overwhelmingly male, disproportionately African-American, and more likely to live alone. The homeless (41 percent of the sample) were less likely to receive food stamps or free food, or to use food pantries. Fewer of them received Medicaid or had health insurance. Forty-seven percent had no income in contrast to 29 percent of the total sample. (*Am J Public Health*. 1991;81:914–917)

Introduction

Soup kitchens, charitable organizations that provide free meals for consumption on the premises, were once regarded as relics of the depression era, useful only to serve the victims of disasters or the disreputable poor in the skid row sections of large cities. Since the early 1980s, however, soup kitchens have again become a significant source of food for the poor.[1]

In New York State, where a recent census identified 189 soup kitchens in New York City (NYC) and 138 in upstate New York and Long Island (Upstate),[2] the State Department of Health provides some funding for soup kitchens through the Supplemental Nutrition Assistance Program[3]; thus the future of the kitchens has become a question for public policy. This report presents data from a survey of persons utilizing soup kitchens in New York City. Upstate kitchens were not surveyed because they had recently been studied.[4]

Methods

The survey of 263 soup kitchen clients from 21 sites in New York City was conducted simultaneously with the food pantry study.[5] A random sample of kitchens was drawn from the Bureau of Nutrition census.[2] Within each site, clients were identified through a random sam-

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From the Department of Nutrition and Food Management, Syracuse University, Syracuse, NY (Bowering and Clancy); and the Department of Sociology, Hunter College, New York, NY (Poppendieck). Address reprint requests to Jean Bowering, PhD, Department of Nutrition and Food Management, 034 Slocum Hall, Syracuse University, Syracuse, NY 13244. This paper, submitted to the Journal May 30, 1990, was revised and accepted for publication January 22, 1991

TABLE 1—Characteristics of Soup Kitchen Clients

	% of Responses				
	Entire Sample (n=263)	Homeless (n=108)	Not Homeless (n=155)	χ^2 Value	p
Gender					
Female	23	13	30		
Male	77	87	70	10.767[a]	0.001
Age of Respondent (years)					
35 or under	57	64	52		
36–59	36	35	37		
60 or older	7	1	11	9.592[b]	0.008
Ethnicity					
African-American	76	69	82	6.216[a]	0.013
Hispanic	16	23	11	6.915[a]	0.009
White	5	6	3	1.520[a]	0.218
Other[c]	3	2	4	—	—
Living Situation					
Alone	62	95	39	85.084[a]	0.0001
With Children	13	0	22	25.250[a]	0.0001
Someone in Household Employed	25	18	29	4.609[a]	0.032
Receive Food Stamps	32	20	40	11.282[a]	0.001
Receive Medicaid	27	15	36	14.411[a]	0.0001
No Medical Insurance	51	79	33	51.432[a]	0.0001

[a] Homeless vs Not Homeless; Chi-square (df=1)
[b] Homeless vs Not Homeless; Chi-square (df=2)
[c] Includes Asians and Native Americans

TABLE 2—Income Sources of Soup Kitchen Clients

	% of Households[b]				
	Entire Sample	Homeless	Not Homeless	χ^2 Value[a]	p
No Income	29	47	16	26.124	0.0001
Wages	22	17	26	3.093	0.079
AFDC (Aid to Families with Dependent Children)	5	1	8	6.929	0.008
Other Public Assistance	9	6	11	2.432	0.119
SSI/Other Disability	8	2	13	10.142	0.001
Social Security	8	3	12	7.543	0.006
Home Relief	7	3	10	5.383	0.020
Other[c]	7	2	11	—	—

[a] Homeless vs Not Homeless; Chi-square (df=1)
[b] Percentages do not add up to 100% because more than one response was possible
[c] Includes Veteran's benefits, Child support/alimony, Retirement/pension, and Unemployment benefits

pling procedure while waiting in line, but the interviews took place either before or after eating with as much privacy as possible. Interviewer selection and training were similar to the food pantry study.[5] The numbers of meals served by each site ranged from a few to 209,000 annually. As with the pantries, large and small sites were appropriately represented in the sample.[5]

Results

Eighty percent of clients had used soup kitchens for less than a year; the remainder were long-term users. Table 1 summarizes other characteristics of soup kitchen users in our sample. Compared with the general population of not homeless persons living below the poverty line in New York City[6] the soup kitchen users we studied were overwhelmingly male, disproportionately African-American, not elderly, and far more likely to live alone. Among those who did live with others, soup kitchen users (both homeless and not homeless) were much less likely than members of the general poverty population to live in households with children.

Compared to the not homeless, the homeless (the 41 percent of our sample who lived in shelters, or in the streets, or "no one place") were more likely to be male, younger, Hispanic, living alone, and unemployed. Fully 47 percent of them had no income.

Twenty-two percent of the entire sample earned wages (Table 2) and there was no significant difference on this between homeless and not homeless. Twenty-nine percent of the sample received welfare benefits, and the homeless were significantly less likely to receive assistance, including SSI and Social Security. The median monthly per capita income (calculated without food stamps) of the entire sample was $120, with 61 percent of them below 50 percent of poverty. The

TABLE 3—Income Distribution and Relation to Poverty Level

	% of Responses				
	Entire Sample (n=243)	Homeless (n=103)	Not Homeless (n=140)	Test Statistic	p
Household Income[a]					
$0	29	47	16		
$1–300	31	38	25		
$301–600	24	9	34		
$601–1000	9	5	13		
$> 1000	7	1	12	50.135[b]	0.0001
Income < 50% of Poverty	61	80	48	25.227[c]	0.0001
Income < 100% of Poverty	83	91	78	7.754[c]	0.005
Median % of Poverty	32%	3%	56%	−6.184[d]	0.0001
Median Monthly per Capita Income					
Excluding Zero Income[e]	$200	$164	$250	−2.192[d]	0.028
Including Zero Income	$120	$ 12	$186	−5.242[d]	0.0001

[a] Income excludes food stamps
[b] Homeless vs Not Homeless; Chi-square (df=4)
[c] Homeless vs Not Homeless; Chi-square (df=1)
[d] Homeless vs. Not Homeless; Wilcoxon Rank Sum Test
[e] Excludes 71 persons with zero income

TABLE 4—Demographic Characteristics of Food Stamp Users (+ FS) and Non-Users (− FS)

	% of Responses							
	Homeless				Not Homeless			
	+ FS (n=22)	− FS (n=86)	χ^2 Value	p	+ FS (n=62)	− FS (n=93)	χ^2 Value	p
Gender								
Female	14	86			40	60		
Male	21	79	0.367[a]	0.545	40	60	0.005[a]	0.943
Age of Respondent (years)								
35 or under	19	81			31	69		
36–59	23	77			49	51		
60 or older	0	100	0.491[b]	0.782	50	50	5.200[b]	0.074
Living/Employment Situation								
Live Alone	19	81	1.245[a]	0.264	51	49	4.906[a]	0.027
With children	0	0	—	—	38	62	0.839[a]	0.360
Someone employed	32	68	1.717[a]	0.190	20	80	10.498[a]	0.001

[a] Food Stamp Users vs Non-Users; Chi-square (df=1)
[b] Food Stamp Users vs Non-Users; Chi-square (df=2)

median monthly income of homeless respondents was $12 (Table 3) and 80 percent had incomes below 50 percent of the poverty level.

Approximately 30 percent of respondents received food stamps. Significantly more of the not homeless participated in the program (Table 1) but, among the homeless, there were no significant demographic differences between food stamp users and non-users (Table 4). In the not-homeless population, significantly fewer recipients lived with others, or in households where someone was employed (Table 4). The average monthly per capita food stamp benefit was significantly higher for the homeless recipients ($73 vs $40, t=−5.087, p=0.0001), undoubtedly reflecting their lower incomes.

Discussion

Comparison of this study with the few recent non-random studies,[7,8] the 1987 study of upstate New York soup kitchen clients,[4] and our food pantry study[5] was done to corroborate the extent of the problems being faced by low income persons. Except for a greater percentage of White clients, data gathered by the Human Resources Administration in 1986 are virtually identical to ours with respect to percent of males, age distribution, homelessness, lack of income, and receipt of public assistance, specifically SS and SSI.[7]

The biggest difference between soup kitchen clients in our NYC sample compared to Upstate clients[4] was the percentage of homeless (41 percent in NYC and 17 percent Upstate). Homelessness in NYC was also reflected in other differences: less use of food stamps, housing subsidies, and health insurance. The NYC clientele was younger, and many fewer received Social Security. More clients Upstate (more of whom were female) were disabled, retired, and received welfare benefits (AFDC, Home Relief) or SSI.

When NYC soup kitchen clients were compared with NYC food pantry users studied at the same time,[5] there were expected differences in age, ethnicity, and homelessness. The educational level was higher among soup kitchen respondents. Significantly fewer soup kitchen clients utilized food stamps or Medicaid, but there were no significant differences in the use of child nutrition and senior feeding programs.

There was a highly significant difference in the percentage of soup kitchen clients who had no income (29 percent) compared to food pantry users (8 percent). Since 88 percent of homeless soup kitchen users received no public assistance, 80 percent received no food stamps, and 79 percent had no health insurance, the precariousness of their lives is evident. A recent report from the National Academy of Sciences documents the fact that homeless people suffer more illnesses and injuries and are at greater risk for a variety of health problems including malnutrition.[9] Soup kitchen clients often eat only once or twice per day and soup kitchen meals tend to be nutritionally inadequate.[10]

Most of the clients in soup kitchens are obviously eligible for public assistance but do not receive it. Many have been unable to meet paperwork requirements; recommendations have been made for reducing these barriers so that more people can receive benefits.[6] However, full public assistance payments will not cover housing costs for most clients. The New York State courts have recently ruled that the Commissioner of Social Services must establish housing allowances that bear a "reasonable relation" to housing costs.[11] For many soup kitchen clients, adequate housing subsidies would provide more funds for food. For others, who are alcohol or drug abusers, other programs are necessary. In the meantime, the need for meals is growing and the size of the problem is too large to be met by the uncoordinated, largely voluntary Emergency Feeding System. ☐

Acknowledgments

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Farm Tractors and Mandatory Roll-Over Protection Retrofits: Potential Costs of the Policy in New York

Timothy W. Kelsey, PhD, and Paul L. Jenkins, MS

ABSTRACT

Tractor roll-overs are the leading cause of fatal farm accidents, accounting for more than one-fourth of all agriculturally related deaths. Most of these deaths could be prevented if the tractors were equipped with roll-over protective structures (ROPS). This study estimates the number of tractors in New York without ROPS, projects their retirement, and then estimates the number of lives which would be saved if ROPS were retrofitted on old tractors. The basic costs associated with mandating ROPS are calculated from these estimates. The minimum economic cost of mandating ROPS is $511,136 per life saved for the retrofits, and an additional $253,254 per life saved for every $1 million spent annually on enforcement. It is concluded that a policy mandating ROPS on all tractors would be expensive, but should be considered with particular attention to the need for and cost effectiveness of enforcement. (*Am J Public Health*. 1991;81:921–923)

Introduction

Agriculture consistently has one of the highest fatal accident rates of any occupation,[1] and tractor roll-overs play a prominent role in such accidents. Tractors are associated with about 50 percent of fatal agricultural accidents,[2] with roll-overs typically accounting for over half of these.[3-6] Evidence from Sweden suggests that most of these roll-over fatalities could have been avoided if the tractors involved had been equipped with roll-over protective structures (ROPS) to prevent operators from being crushed by an overturn.[7]

Most farm tractors in the United States lack ROPS, with only 19 percent of tractors in Pennsylvania[6] and 15 percent in one New York county* so equipped. ROPS became available in the 1960s, but did not become standard equipment on many tractor models until the late 1970s and early 1980s. In 1985, manufacturers voluntarily agreed to make them standard equipment on all new tractors.

Various proposals have been made to make ROPS mandatory in the US.[4,8] This study analyzes the National Coalition for Agricultural Safety and Health (NCASH) proposal to offer farmers retrofit incentives for five years and require ROPS on all tractors after 10 years. The potential economic cost per life saved from such a policy in New York is estimated using cost-effectiveness analysis, a common approach to appraising public health policies.[9-15] The cost shifting effects of incentives are not addressed in this analysis because of the study's economic (as opposed to financial) focus.

Methods

There are few available data on the age distribution of tractors currently on US farms or their ROPS status. The magnitude of the effect of ROPS in reducing the number and severity of non-fatal injuries is uncertain, especially when the operator is not wearing a seat belt, although anecdotal evidence suggests almost all deaths are prevented.

For this study the potential number of tractors which would be retrofitted in New York State was estimated by interpolating from US Census of Agriculture, Farm and Industrial Equipment Institute, and Pennsylvania tractor data, using the 4 percent retirement rate implied by data in the Census of Agriculture. There are no available data about the percentage of tractors of different ages in New York without ROPS, so the results of a survey in Pennsylvania by Huizinga and Murphy[6] had to be used. The geography, agricultural production, and economic situation of farmers in Pennsylvania and New York are roughly similar, making their data a relatively good approximation for tractors in New York. A total of 59,873 tractors would need to be retrofitted according to these calculations. It was assumed that all tractors covered by the 10-year requirement would be retrofitted during the first five years of the program, when the incentives are in effect (see Figure 1).

An average of nine people a year die from tractor roll-overs in New York State.** The number of deaths which would be averted by the retrofit policy was estimated by combining this with the current percentage of tractors without ROPS and the estimated pace of retrofits, under

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From the New York Center for Agricultural Medicine and Health. Address reprint requests to Timothy W. Kelsey, PhD, Agricultural Economist, New York Center for Agricultural Medicine and Health, One Atwell Road, Cooperstown, NY 13326. Mr Jenkins is a statistician with the Center. This paper, submitted to the Journal April 17, 1990, was revised and accepted for publication January 23, 1991.

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*Unpublished study of Otsego County, New York. Cooperstown, New York: New York Center for Agricultural Medicine and Health.

**Unpublished data on New York agriculturally related accidents. Cooperstown, NY: New York Center for Agricultural Medicine and Health.

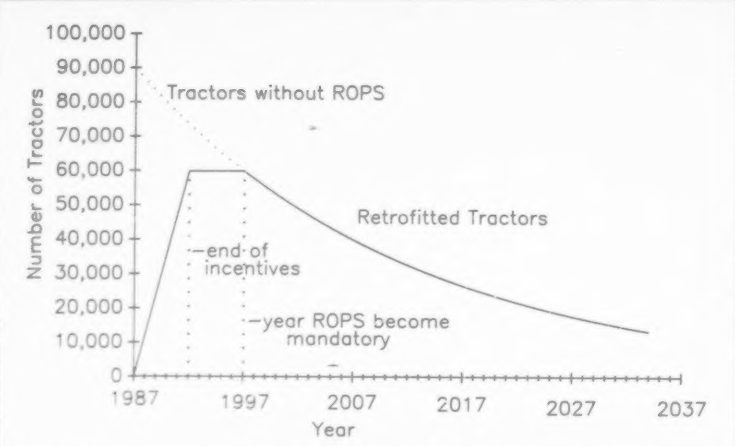

FIGURE 1—Estimated Number of Tractors Originally Manufactured Without ROPS Operating in New York if ROPS Became Mandatory on all New Tractors

TABLE 1—Cost Per Life Saved

	Cost Per Life Saved*	
Life-Saving Effectiveness of ROPS	For ROPS	For Every $1 Million Spent Annually on Enforcement
75 Percent	$684,729	$337,672
100 Percent	$511,136	$253,254

*Discount rate = 5 percent

two alternative assumptions about the life-saving effectiveness of ROPS; 100 percent and 75 percent.

The costs of such a policy would include enforcement, as well as purchase and installation of the ROPS. Without enforcement there is no guarantee that all tractors will be retrofitted or that retrofitted ROPS will not later be removed. In Pennsylvania, 38 percent of the tractors manufactured since 1986 do not have ROPS, even though beginning that year tractors without ROPS were sold only under special circumstances. This occurred in part because farmers removed the ROPS after purchasing the tractor.[6] The reasons for removals have not been comprehensively studied, although the inability to operate a ROPS equipped tractor in buildings with low ceilings is a common explanation. The costs of enforcement can be significant, but because the NCASH proposal lacks an enforcement plan such costs will not be directly integrated into the analysis.

The specific cost of retrofitting tractors was determined by calling five ROPS manufacturers and wholesalers, including John Deere, Inc, and Ford/New Holland. The contacts indicated that increased sales would not affect most prices. The average uninstalled retail price reported for most ROPS was $632, which means the average final per unit cost will be approximately $700 with installation. The present value of all the retrofits was calculated assuming that the retrofits occur evenly throughout the five-year incentive period. Both costs and the lives saved were discounted with a 5 percent discount rate.

Results

A policy mandating ROPS in 10 years requiring any tractor in New York to have ROPS to be operated and providing economic incentives for retrofits during the first five years of the program could have a total ROPS cost of approximately $36.3 million. The present value of lives saved would be 53 or 71, depending upon the life-saving effectiveness of ROPS. The costs per life saved under the different scenarios are in Table 1. If perfect compliance does not occur and fewer lives are saved, the enforcement cost per life saved would be higher.

Discussion

These estimates of cost do not include consideration of the effectiveness and cost of enforcement. In addition, the model of tractor retirements biases the number of lives saved upwards. Another upward bias in estimation of the lives saved is that we could not include the natural reduction in farmers' exposure to risk over the years as tractors are slowly shifted from field work to more stationary roles and their hours of use decrease. Finally, non-pecuniary aspects of requiring ROPS retrofits are not included in these estimates.

Based on the projected cost without enforcement, ROPS would be similar to other life-saving interventions, but the need for enforcement could drastically increase cost. One reviewer found costs less than $1 million for auto safety standards and greater than or equal to $3 million for most efforts to regulate toxic substances.[16] Graham and Vaupel reviewed 35 studies of life-saving programs, and found the cost per life saved ranged from no cost for compulsory seat belt use, $60,000 for mandatory smoke detectors, $255,000 for the 1966 Motor Vehicle Safety Act, $408,000 for mandatory air bags, and $1,200,000 for the 55 mph speed limit, to over $3,500,000 for policies regarding toxic substances.[17]

Mandating ROPS retrofits is worth considering because of its life-saving impact, but enforcement must be directly and carefully debated because the policy impact will depend so much upon it and because enforcement could easily be the largest expense. In any event, a retrofit policy will not be cheap. ☐

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The Use of Direct Mail to Increase Clinician Knowledge: An Intervention Study

Donald Sadowsky, DDS, MPH, PhD, and Carol Kunzel, PhD

ABSTRACT

A probability sample of American general practitioner dentists, 40 years of age or older, in solo private practice, was the target of two direct mail interventions offered at two different times, to test whether knowledge regarding prophylaxis of patients at risk for infective endocarditis could be improved. Tests of knowledge were responses in a mail questionnaire to clinical vignettes, designed to elicit the content of antibiotic regimens used for patients at risk. The research design enabled detection of (1) the effect of the interventions; (2) the differences in their effect; (3) the attenuation of their effect; and (4) the effect of time. Where baseline knowledge was low, it was improved and did not rapidly disappear. Both mail interventions were equally effective, in most instances, and there was no attenuation of the interventions' effect and no effect of time on the control groups' knowledge over a 4½ month period. The results suggest that it is possible to improve clinicians' knowledge of expert recommendations through direct mail intervention. (*Am J Public Health.* 1991;81:923–925)

Introduction

Direct mail has been used relatively infrequently, except by the pharmaceutical industry, to "market" ideas to clinicians. The results of several studies[1-7] testing the efficacy of mailed information in improving clinicians' knowledge are equivocal. Moreover, the effects of these efforts are often difficult or impossible to discern because of the absence of a sound experimental design.

Since the mid-1950s, the American Heart Association (AHA) has been trying to inform physicians and dentists about the details of a variety of antibiotic regimens that have been determined to be efficacious in the prevention of infective endocarditis, depending on the risk factor and the procedure to be performed. As the state of the art has changed, these AHA recommendations have been revised.[8-13] Dentists have been a target group for this AHA program because any procedure that produces mucosal bleeding may produce a transient bacteremia which, in turn, may seed a heart valve previously scarred by, for example, rheumatic heart disease. The ultimate sequela of this series of events may be infective endocarditis.

Our project had four aims: (1) to determine whether simple inexpensive mail interventions could be designed to improve the knowledge of dental general practitioners (GPs) about the management of patients at risk for infective endocarditis; (2) to design two interventions, one with a theoretical component and one without it, to determine whether the additional material would have an additional effect; (3) to determine whether the effects of the interventions, if any, would be maintained over time; and (4) to determine whether clinicians' knowledge would change over time, independent of the planned interventions.

Methods

After a probability sample of 1148 dentists, classified as GPs, 40 years or older, in private practice within the continental United States, was drawn from

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Donald Sadowsky and Carol Kunzel are with the Albert Einstein College of Medicine. Dr. Sadowsky is Professor, both in the Department of Dentistry and in the Department of Epidemiology and Social Medicine; Dr. Kunzel is Assistant Professor in the same departments.

Requests for reprints should be sent to Dr. Donald Sadowsky, Albert Einstein College of Medicine, Dental Research, Van Etten Hospital, Room 5C15, Bronx, NY 10461.

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TIME	C1	C2	E1	E2	E3	E4
TIME^{+} ZERO T_0 No measurement for any group at this time	No Stimulus	No Stimulus	Receives Recs (Intervention 1)	Receives Recs (Intervention 1)	Receives Recs + Theory (Intervention 2)	Receives Recs + Theory (Intervention 2)
TIME ONE T_1 (4 1/2 Mos. Later)	Receives DCI*		Receives DCI* $\overline{E1} - \overline{C1}$ Equals Effect of Recs		Receives DCI* $\overline{E3} - \overline{C1}$ Equals Effect of Recs + Theory $\overline{E3} - \overline{E1}$ Equals Effect of Theory	
TIME TWO T_2 (Nine Mos. Later)		Receives DCI* $\overline{C2}_{T2} - \overline{C1}_{T1}$ Test for "Natural History" Effect		Receives DCI* $\overline{E1}_{T1} - \overline{E2}_{T2}$ Test for Attenuation of Recs		Receives DCI* $\overline{E3}_{T1} - \overline{E4}_{T2}$ Test for Attenuation of Recs + Theory

*Data Collection Instrument

$\overline{C1}$, $\overline{C2}$, $\overline{E1}$, $\overline{E2}$, $\overline{E3}$, $\overline{E4}$ = Mean Knowledge Level of Group

$^{+}$ Knowledge Level of All Six Groups Assumed Equal at Time Zero

FIGURE 1—Groups (Each group is representative of solo US GPs, 40 years or older and in private practice)

the American Dental Association's master list containing names of all dentists in the United States, the list was cleaned and criteria for eligibility were ascertained via postcards. Remaining potential respondents (686 GPs, 40 years of age and older, in solo private practice*) were randomly assigned to one of six experimental conditions. An overall response rate of 87.20% was achieved.

Four groups were contacted at baseline time zero (Fig 1). Two of these groups received a laminated placard (intervention 1), containing AHA information.[17] The reverse side of the placard was blank. The other two groups received the same placard but with 11 statements on the other side providing information about the theoretical underpinning of the AHA recommendations (intervention 2) with the expectation that enhancing understanding might help increase knowledge.[16] All placards were mailed with a cover letter encouraging the ongoing use of the placard as a resource in the office setting. Knowledge is used in this article to indicate availability of information. The remaining two groups (controls) were not contacted at time zero.

At time 1, 4½ months later, members of one control group, one group who had received intervention 1, and one group who had received intervention 2 were mailed an 8-page data collection instrument, along with a cover letter and a check for 10 dollars. At time 2, 4½ months after time 1, the same procedures were repeated for a second control group, for another group who had received intervention 1, and for a second group who had received intervention 2.

A precoded, forced choice, data collection instrument used varieties of clinical vignettes to elicit information. Dentists were asked to indicate the content of the antibiotic prophylactic regimen they might prescribe or administer for patients at different types of risk. Answers to questions about different patient conditions were assessed as "correct" = 1; "incorrect, unsure" = 0; "blank" = missing, using the standard of the most recent (at the time of the study) AHA recommendations. The resultant 0–1 dichotomous variables were analyzed as a special case of proportions, using either three-group or two-group χ^2 tests. The Yates correction was used for all two-group comparisons. SPSS/PC was used to perform all statistical computations.

Results

There was no significant attenuation in knowledge from one time period to the next in groups receiving the interventions, and knowledge scores remained stable over time in the control groups.

Because analyses indicated that time itself had no effect on knowledge scores, data for both control groups, for both intervention 1 groups, and for both intervention 2 groups were combined to ascertain the net effect of the interventions. χ^2 analyses contrasted the proportion of correct answers in the three types of experimental groups. Results indicate that, with the exception of the valvular heart disease patient, experimental group membership had a statistically significant impact on the proportion of correct answers about regimen components, as well as a statistically significant impact on correct answers about bacteremia duration, an important theoretical variable underlying the recommendations (Table 1).

There were no statistically significant differences between the two interventions for knowledge about the prosthetic heart disease patient, the child valvular heart disease patient, and the prosthetic heart valve patient with a penicillin allergy. Intervention 1 did not affect correct knowledge of managing the patient on a continuing antibiotic regimen and of bacteremia duration. However, intervention 2 had a statistically significant impact on both these knowledge items.

Discussion

The relative absence of an effect from the addition of scientific justifications for the AHA recommendations (intervention 2) can perhaps be explained by the findings of a comprehensive study by the RAND Corporation,[18] i.e., that consensus statements "are more likely to be effective if they are succinct, didactic, and oriented toward a clinical rather than a scholarly audience." The RAND study also provides a likely explanation for the uneven effects of the interventions: the negligible effect of the interventions on knowledge about the management of patients with a history of valvular heart disease is probably accounted for by the relatively high knowledge about antibiotic management of patients with this particular risk factor before the interventions. In reverse, this may also be a partial explanation for why knowledge about both the management of children and about patients with prosthetic heart valves improved considerably: for these conditions, baseline knowledge was very low and there was an opportunity for a great deal of improvement.

The improved knowledge scores of groups receiving intervention 2, on items about bacteremia duration and the man-

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*Eighty percent of American dentists are GPs[14] and, on a test of knowledge of endocarditis prevention, they scored dramatically lower than a cohort of dental specialists.[15] Also, dentists in solo general practice, 40 years or older, are most likely to be deficient in knowledge about the proper management of patients at risk for infective endocarditis.[16]

TABLE 1—Net Effect of Interventions (Time 1 and Time 2 Groups Combined)

Type of Patient and Knowledge Item Tested	Percentage of Correct Answers, %(n)				
	Controls	Intervention 1	Intervention 2	χ^2 df = 2	2-Tailed P*
VHD[a] patient					
Preprocedure drug	85 (149)	91 (169)	84 (156)	4.319	.115
Preprocedure dose	59 (110)	69 (135)	63 (118)	3.976	.137
When to begin Rx	68 (127)	72 (144)	73 (138)	1.502	.472
PHV[b] patient					
Preprocedure drug	15 (26)	28 (49)	23 (40)	8.329	.016
Preprocedure dose	14 (25)	25 (48)	26 (47)	9.339	.009
When to begin Rx	17 (32)	29 (58)	25 (46)	7.855	.020
80 lb VHD[a] Child					
Preprocedure dose	28 (53)	41 (82)	44 (85)	11.620	.003
Patient on continuing regimen of antibiotics					
Alteration in Antibiotic regimen required	32 (58)	32 (65)	43 (82)	7.225	.027
PHV[b] allergic patient					
Preprocedure drug	10 (19)	22 (43)	19 (36)	9.640	.008
Bacteremia duration	22 (41)	22 (44)	38 (74)	17.901	.000

[a] Valvular heart disease.
[b] Prosthetic Heart Valve.
*2-tailed test was used because direction of change among three groups could not be anticipated.

agement of patients on a continuing antibiotic regimen, can be explained by the theoretical nature both of these items and of the intervention.

The experiment described provides evidence to support the notion that it is possible to improve clinicians' knowledge through direct mail intervention and that the effect does not rapidly disappear. More work must be done before the conditions under which this can be accomplished will be specified with any precision. Our study may provide some helpful hints about what works and what does not work. ☐

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Service Needs of Severely Disturbed Children

ABSTRACT

Background: As one piece of a statewide Children's Mental Health System Analysis conducted in Washington State, the specific service needs of severely emotionally disturbed children were investigated.

Methods: The primary caseworkers of 3398 children under the care of the state responded to a survey of the child's treatment history, social and clinical condition, and service needs. Service need data on the 2455 children in the sample who were classifiable as severely emotionally disturbed were examined using log linear analysis.

Results: School- and home-based services were frequently reported needs of school-aged children. Adolescent children were reported to be in need of services which prepare them for adulthood. Services endorsed for female children tended to focus on assistance to family members, therapy, and protection, whereas services endorsed for male children focused on home removal, substance abuse treatment, and preparation for employment.

Conclusions: Severely emotionally disturbed children are in need of a variety of nontraditional mental health services. Significant differences in the needs of male versus female children, and of differently aged children, are present that have serious implications for service system development. (*Am J Public Health.* 1991;81:975–980)

Eric W. Trupin, PhD, Amy Forsyth-Stephens, MSW, and Benson P. Low, PhD

Introduction

The treatment needs of severely and chronically mentally ill individuals, of any age, vary considerably depending upon personal characteristics (e.g., degree of disability, nature of symptomatology, willingness to accept treatment), and situational or environmental characteristics (e.g., availability of natural supports, residential situation, mobility of the individual). Ideally, the range of available services would be so broad, and care givers so flexible, that all clients would receive an appropriate, individualized menu of services. Unfortunately, human services, especially mental health, are largely program-driven, with clients matched to programs, rather than services matched to clients. Given that this infrastructure of care provision may be difficult to overturn, it is imperative that the development and design of state and local programs be based upon a concrete knowledge of client needs.

Services for children* are especially subject to "blind" planning and implementation. There is a dearth of research in the area of children's mental health, most notable on the topic of severely mentally ill children.[1] Analogous to the study of chronic mental illness among adults, the area of severe emotional disturbance (SED) among children is a relatively recently established field of research. Attention was drawn to the issue in 1982, when the unsettling book *Unclaimed Children*[2] reported that two-thirds of all SED children do not receive the services they need.

The book went on to document the failure of federal and state governments to acknowledge and attempt to meet the mental health needs of children, resulting in a vast and deeply troubled population with no advocacy, options, or hope. Five years later, the Office of Technology Assessment of the United States Congress[3] underscored accusations put forth in *Unclaimed Children.*

In 1984, the National Institute of Mental Health initiated the Child and Adolescent Service System Program (CASSP) in an effort to address the national crisis in children's mental health.[4] The goal of CASSP was, and continues to be, the promotion of system development through the funding of 40 state-level planning initiatives and two national research centers for the study of SED among children.

Most of the research to date on service needs of SED children has been produced by state human service departments in response to legislative mandates or class action suits, or by CASSP-funded research institutes. Based upon work completed by many states, and by the CASSP research centers, an assumption has been promulgated that a complete range of services, on a continuum from most to least restrictive, must be developed to adequately meet the needs of severely emotionally disturbed children.[5] Most recently, the "continuum of care" idea has been replaced by the concept of a "system of care" for SED children,[6]

From the University of Washington School of Medicine, Department of Psychiatry & Behavioral Sciences. Address reprint requests to Eric W. Trupin, PhD, Division of Community Psychiatry, Department of Psychiatry & Behavioral Sciences, School of Medicine, University of Washington, Mailstop RP-10, Seattle, WA 98195. This paper, submitted to the Journal July 9, 1990, was revised and accepted for publication January 23, 1991.

*In this article, "Children" refers to anyone under the age of 19 years.

which has a somewhat broader connotation. With the system perspective, more than a comprehensive spectrum of appropriate services is necessary. Just as critical are the underlying facilitating mechanisms, structures, and administrative processes to ensure that services are delivered and received in a coordinated, child-focused manner.

There have been no empirical studies of the diverse service needs of severely emotionally disturbed male children versus female children. Similarly,there have been no reports of service need data by developmental groupings (i.e. young children, preteens, and adolescents). Only one report has documented the age and sex breakdowns of an SED child population. In North Carolina, Behar[7] reports that the population of children qualifying as SED was 80 percent male, 20 percent female. Over three-fourths (78 percent) of the SED children were between the ages of 12 and 17; 14 percent were under the age of 12.

This paper presents the findings of a statewide survey of the specific service needs of 1617 severely emotionally disturbed children. The large sample size of the survey permits analysis of service need data by age and sex categories. The possibility of identifying distinct ideal service continuums for male and female SED children of various age groups is discussed.

Method

Design

The survey of service needs of SED children took place in the context of a broader study of the children's mental health system in Washington State. This larger project, entitled the Children's Mental Health System Analysis, involved the administration of a Child Mental Health Survey to adult care takers of randomly selected, nonidentified children served by the Department of Social and Health Services (DSHS).

The analysis of data occurred in two stages. The first stage involved the identification of children in the sample who were classifiable as SED (this procedure is described below under "Classification of Children as SED"). The second stage involved the examination of reported service need data for the subjects who were classified as SED according to our model.

A significant design feature was the nonidentified status of children surveyed. Use of identified data would have required informed consent for every child in the survey. Apart from the logistic demand such a requirement would have brought, institutional or agency refusals would have become a significant problem. While the use of nonidentified data facilitated more comprehensive surveying of children, it did limit control over the subject selection process, and precluded empirical analysis of subject attrition.

Subjects

The subjects were 3,398 children and adolescents ages 0 to 18 who were selected randomly from the billing rosters of child-serving programs within the State's Department of Social and Health Services (DSHS). Nearly half of the subjects (N = 1,640) were clients of the DSHS Division of Children and Family Services (DCFS), which offers an array of child welfare services including child protective services, foster care, group home care, in-home intervention, and emergency out-of-home placement. The remainder of the sample was made up of clients of the DSHS Division of Juvenile Rehabilitation (DJR; institutions, group homes, and parolees; N = 1,035); the DSHS Mental Health Division (DMH; inpatient, residential, day treatment, and outpatient; N = 524); and the DSHS Division of Developmental Disabilities (DDD; institutions and group homes; N = 199).

Sample sizes were calculated by taking into consideration the population size, the *a priori* estimate of the rate at which SED would occur (50 percent in all DSHS groups), the range of the intended confidence interval (+5 percent), and the probability level for the confidence interval (95 percent). Sample sizes were inflated by a factor of 15 percent, to compensate for subject attrition. As an exception, the DSHS Division of Juvenile Rehabilitation made an administrative decision to survey all clients in their programs. All clients in programs with caseloads of fewer than 150 were surveyed.

Respondents

The survey respondents were 3,398 adult DSHS professionals or staff members of agencies under contract with DSHS. Two-thirds of the respondents (65 percent) were case workers or facility staff employed either directly by DSHS, or by a contracted agency. The remaining one-third was made up of probation/parole officers and independent mental health professionals. Ninety-two percent of the respondents had at least a four-year college degree. The mean length of time that a respondent had known the surveyed child was 12.2 months.

Survey Instrument Development and Reliability

The data collection instrument, entitled the Child Mental Health Survey,* was developed after review of materials used by other states in similar but smaller-scale efforts.[8] The survey instrument contained six sections with questions pertaining to: 1) sociodemographic background; 2) exposure to 13 mental health risk conditions; 3) current behavioral adjustment as measured by the Child Behavior Checklist (CBCL);[9] 4) past mental health treatment and placement history; 5) other handicapping conditions; and 6) mental health and social service needs. The section on service needs required the respondent to identify the subject's past receipt, current receipt, and current unmet need for any of 32 human services.

The Child Mental Health Survey was administered to pairs of teachers of 26 special education pupils and 30 regular education pupils, and to pairs of case workers of 10 children placed in mental health treatment programs in the community. Each of the respondents independently completed surveys on the selected child. Reliability coefficients were calculated on four sections of the survey: exposure to mental health risk factors, mental health treatment history, other handicaps, and current behavioral adjustment. The Kappas obtained for the sections on risk factors, treatment history, and other handicaps were in the acceptable and expected range, with most values above .60, and some exceeding .80.[10]

Reliability coefficients were not calculated on service need data. Interrater reliability was less of a concern for the service need data, as the caseworkers who completed the survey forms were highly face valid, as they are the "natural" raters who drive the system of referrals for these SED children. Moreover, a meta-analysis on concordance of raters of children underscores that data from a single respondent typically provide a reasonable sample of what would be provided by other informants of the same type who see the child under generally similar conditions.[10]

Classification of Children as SED

SED children from the sample were identified via MANOVA techniques.

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*The Child Mental Health Survey is available from the first author upon request.

Each subject was treated as a group of one and then compared via Hotelling's T^2 against a criterion SED group. The criterion SED group was defined as all children being served in the State's psychiatric residential treatment centers, and a specialized foster care program for previously psychiatrically hospitalized children. These programs provide the most comprehensive, intensive, and expensive mental health treatment offered by the state mental health system for children. It was the research team's consensus that children meeting the numerous requirements for admission to these programs are inarguably severely emotionally disturbed, and therefore comprise the most valid, conservative, and defensible criterion group for this study.

Each subject was evaluated individually as to the likelihood of being drawn from the criterion group. If the subject was discriminable, the subject was not classified as SED. The independent variables used in the discriminant analysis were: 1) *Mental Health Risk,* scored as the number of risk factors endorsed positively for the subject; 2) *Chronicity,* scored as positive if the subject was rated as experiencing severe aggressiveness and/or withdrawal for a minimum duration of one year; 3) *Service Need,* scored as the number of mental heath services endorsed as "needed" or "currently received"; 4) *Internalizing Behavior Problem,* which was the T-score from the CBCL Internalization Scale; and 5) *Externalizing Behavior Problem,* which was the T-score from the CBCL Externalization Scale. These five measures are widely used and face valid with respect to the detection of the presence of an emotional or behavioral problem in children.[11–14]

Service Need Data Analysis

Service need data were analyzed for those children classified as SED according to the above model. Data on the services indicated as "needed" by SED children were weighted by DSHS population estimates of SED children across age and sex categories. This was done in order to control for the varying age and sex distributions of the four subsamples (i.e. DCFS, DJR, DDD, DMH). The population estimates were based upon the prevalence distribution of SED in our subsamples. A log linear analysis technique was used to test for significant differences in service need among three age groups of SED children (6–11, 12–16, and 17–18 years), and between male and female SED children.

Results

Survey Return Rates

The overall DSHS survey return rate was 75 percent (i.e. 4,516 surveys were distributed, 3,398 were completed and returned). Return rates for the various participating divisions of DSHS ranged from a high of 95 percent for DJR to a low of 66 percent for DCFS. The high overall survey return rate was primarily a function of consistent DSHS involvement with the design and execution of the study. Most supervisors at local DSHS offices required 100 percent survey completion by their caseworkers, and actively monitored this assignment. Additionally, researchers distributed two reminder notices to DSHS supervisors during the time frame allotted for survey completion. All respondents were assured of, and received, feedback on the outcome of the survey.

Service Needs of SED Children

Of our sample of 3,398 children, 2,455 were classified as SED according to our model (described above, $\alpha = .10$). Of these 2,455 SED children identified from our original sample, data on 1,617 were used in this analysis of service needs. A total of 838 cases were eliminated due to the subject's age (less than 6 years) and/or an incomplete survey form. Children ages 0 through 5 were not included in this analysis due to the small size of that subsample ($n = 96$). Table 1 reports the estimated percent of the DSHS SED population needing various mental health and social services.

Across all age and gender categories, over one-half of all SED children needed outpatient mental health treatment (defined as individual, group, or family therapy). Nearly one-half of SED children's families were in need of family support services for the parents and/or siblings. Interagency case management was needed by nearly one-third of SED children statewide. Other services which were commonly needed included school-based mental health services and psychological testing. Of all residential/inpatient services, foster home placement and group home placement were the most commonly reported service needs.

Age and Gender Service Need Differentials

Fourteen of the 32 possible mental health and social services listed on the Child Mental Health Survey showed significant differences in reported need across the three age groups. These are displayed in Table 2.

School children from age 6 through 11 (Group 1) were significantly higher than all other age groups in the need for respite care for their family members (11.0 percent, $p < .000$), and for therapeutic camp services (17.7 percent, $p < .000$).

Children in Age Groups 1 and 2 (ages 6–11 and 12–16) displayed a higher need for school and home-based services than older SED children, specifically school-based mental health services (Group 1: 30.0 percent; Group 2: 27.4 percent), supervised after-school activities (Group 1: 20.3 percent; Group 2: 18.0 percent), and intensive in-home therapy (Group 1: 15.6 percent; Group 2: 18.8 percent).

Children ages 12 through 16 (Group 2) were reported as more in need of foster home placements (18.9 percent, $p < .001$) and group home placements (12.6 percent, $p < .000$) than children in other age brackets. Data show that the need for both of these services peaks in the years 12 through 16, although the need pattern for the two services is quite different across the full developmental range. The need for foster home placement is high (12.3 percent) even within the 6 to 11 year old age group, and grows in the 12 to 16 year old group. At that point the need declines to 11.1 percent in the 17 to 18 year old age group. Group home placement, however, is low for the 6 to 11 year old age group, and then rises dramatically for the 12 to 16 year olds, staying relatively high (11.3 percent) for the 17 to 18 year old group.

Older adolescents (Group 3, ages 17–18) demonstrated a significantly higher need for outpatient alcohol or drug treatment (31.4 percent, $p < .000$), and inpatient alcohol or drug treatment (15.6 percent, $p < .000$) than the other age groups. They also demonstrated a much greater need than other age groups for training in independent living skills (58.6 percent, $p < .000$), vocational services (49.2 percent, $p < .000$), and supervised independent living arrangements (28.5 percent, $p < .000$). For each of these services, the expressed need was nonexistent or extremely low for younger children (ages 6–11), increased for children ages 12 through 16, and peaked for children 17–18 years of age.

Significant sex differences in endorsement of service need were identified among 11 of the 32 service types contained on the Child Mental Health Survey. These differences are also reported in Table 2.

Females were significantly higher than males in their reported need for out-

TABLE 1—Estimated Percent of Severely Emotionally Disturbed (SED) Child Population Needing Service, by Age and Gender

Service	Male Female Total	Ages 6–11 (N = 145) (N = 117) (N = 262)	Ages 12–16 (N = 594) (N = 284) (N = 878)	Ages 17–18 (N = 406) (N = 71) (N = 477)	All Ages (N = 1145) (N = 472) (N = 1617)
Outpatient Services					
Outpatient therapy	M:	60.5	49.3	32.1	49.6
	F:	72.1	63.0	44.9	63.2
	T:	66.0	55.9	37.3	56.0
Case management	M:	14.9	34.5	9.0	25.4
	F:	30.6	38.4	36.2	35.9
	T:	22.2	36.3	19.5	30.3
School-based mental health services	M:	33.9	27.5	9.9	28.4
	F:	25.5	27.4	22.3	26.3
	T:	30.0	27.4	14.9	26.3
Psychological testing	M:	29.1	27.3	22.8	27.1
	F:	15.7	24.4	15.9	21.1
	T:	22.9	25.9	20.0	24.3
Intensive in-home therapy	M:	15.2	17.0	9.8	15.4
	F:	16.1	20.9	15.9	19.0
	T:	15.6	18.8	12.2	17.1
Psychiatric evaluation	M:	15.8	20.4	16.7	18.6
	F:	14.3	14.5	10.6	14.0
	T:	15.1	17.6	14.3	16.5
Outpatient alcohol/drug treatment	M:	2.6	18.9	40.0	17.9
	F:	2.1	14.5	18.7	11.7
	T:	2.4	16.8	31.4	15.0
Day treatment	M:	11.9	9.1	7.1	9.5
	F:	4.9	6.1	7.7	6.0
	T:	8.6	7.7	7.3	7.9
Therapeutic preschool	M:	6.0	1.4	2.0	2.7
	F:	4.5	0.9	2.4	2.0
	T:	5.3	1.1	2.1	2.4
Medication evaluation	M:	3.4	3.3	0.7	2.9
	F:	2.3	1.7	5.4	2.3
	T:	2.9	2.5	2.6	2.6
Residential or Inpatient Services					
Foster home placement	M:	11.1	16.1	2.4	12.7
	F:	13.7	22.0	23.9	20.1
	T:	12.3	18.9	11.1	16.1
Group home placement	M:	3.5	15.3	13.3	11.9
	F:	3.8	9.7	8.3	6.0
	T:	3.6	12.6	11.3	10.1
Inpatient alcohol/drug treatment	M:	0.0	8.4	25.5	8.9
	F:	0.0	8.2	0.8	5.2
	T:	0.0	8.3	15.6	7.2
Crisis residential placement	M:	0.0	3.5	1.4	2.0
	F:	4.3	9.5	0.2	7.0
	T:	2.0	6.4	0.9	4.5
Psychiatric residential treatment	M:	2.2	2.7	2.8	2.6
	F:	1.0	2.9	5.1	2.7
	T:	1.6	2.6	3.8	2.6
Psychiatric hospitalization	M:	0.0	1.9	0.7	1.2
	F:	0.0	1.9	0.0	1.2
	T:	0.0	1.9	0.4	1.2
Services to the Family					
Family support group	M:	44.6	46.4	38.0	44.6
	F:	43.3	55.5	50.8	51.7
	T:	44.0	50.8	43.2	47.9
Respite care	M:	11.1	3.5	1.9	5.2
	F:	10.8	7.7	5.1	8.2
	T:	11.0	5.5	3.2	6.6
Services to Adolescents					
Training in independent living skills	M:	0.0	20.9	58.4	21.3
	F:	1.1	20.5	59.0	20.0
	T:	0.5	20.7	58.6	20.7
Vocational services	M:	0.0	18.5	56.3	19.5
	F:	1.1	16.2	38.6	14.9
	T:	0.5	17.4	49.2	17.4

continued

patient therapy (61.7 percent, p < .000), foster home placements (19.1 percent, p < .001), respite care (8.0 percent, p < .000), supervised independent living arrangements (7.6 percent, p < .001), and child protective services referral/intervention (9.6 percent, p < .000).

Males were rated as having a significantly greater need for outpatient alcohol or drug treatment (16.2 percent, p < .000), inpatient alcohol or drug treatment (7.9 percent, p < .000), group home placement (10.6 percent, p < .000), training in independent living skills (19.0 percent, p < .000), and supervised afterschool activities (19.8 percent, p < .006).

Discussion

Our survey of the professional caregivers of SED children generated a useful data base on the service needs of this population. While previous research and program planning efforts have identified only general categories of services appropriate for SED children, this broad-based study is the first to produce quantitative data on the extent of the SED population reported to be in need of a wide range of mental health, health, social, and support services. It is also the first study that further defines age and gender differences in types of needed services.

Overall Patterns of Service Need

Across age and sex groupings, the most frequently reported needs, in descending order, were for: 1) outpatient therapy; 2) family support groups; 3) intersystem case management; 4) school-based mental health services; 5) psychological testing; 6) training in independent living skills; 7) intensive in-home therapy; and 8) supervised after-school activities. With the exception of outpatient therapy and psychological testing, all of these services would be considered "non-traditional" mental health services, or even outside the realm of mental health altogether.[6]

We had anticipated that our survey of professionals would result in a high expressed need for residential or inpatient services, since respondents' level of familiarity with other more progressive treatment approaches was questionable. We reasoned that direct care providers might be hesitant to endorse non-traditional services as needed, especially if such services were not accessible to them.

Clearly, this was not the outcome. The strong presence of CASSP in Washington State may be credited for a good

TABLE 1—Continued

Service	Male Female Total	Ages 6–11 (N = 145) (N = 117) (N = 262)	Ages 12–16 (N = 594) (N = 284) (N = 878)	Ages 17–18 (N = 406) (N = 71) (N = 477)	All Ages (N = 1145) (N = 472) (N = 1617)
Supervised independent living arrangement	M:	1.0	5.1	27.1	7.5
	F:	0.0	7.9	30.5	8.5
	T:	0.6	6.4	28.5	7.9
Other Services					
Supervised afterschool activities	M:	22.4	22.8	10.7	20.8
	F:	17.9	12.8	2.3	12.9
	T:	20.3	18.0	7.4	17.1
Transportation to needed services	M:	22.7	13.9	5.1	14.8
	F:	12.4	15.0	5.5	13.1
	T:	17.9	14.4	5.3	14.0
Recreational services	M:	13.7	20.1	7.3	16.4
	F:	10.4	11.4	2.9	10.1
	T:	12.2	15.9	5.5	13.5
Therapeutic camp	M:	20.0	10.9	2.6	12.0
	F:	15.0	10.0	4.9	10.7
	T:	17.7	10.5	3.5	11.4
Special education	M:	15.3	12.4	12.7	13.2
	F:	6.3	10.3	3.0	8.4
	T:	11.1	11.4	8.8	11.0
Child protective services	M:	5.5	2.3	1.2	3.0
	F:	11.5	6.0	6.7	7.5
	T:	8.3	4.1	3.4	5.1
In-home education	M:	5.6	4.2	1.3	4.1
	F:	5.5	4.1	10.2	5.2
	T:	5.6	4.1	4.9	4.6

deal of the progressive thinking of the state's professional care givers through conferences, newsletters, and an intensive effort to begin the creation of local intersystem case coordination committees. Not recognizing the significant impact of these educational efforts, we underestimated the knowledge base of our respondent pool.

Unfortunately, Washington State's current service system reflects the national reality, in that treatment resources continue to be focused on a proportionally small number of children who receive care on inpatient psychiatric units or other intensive residential treatment programs.[3,15] Our data support the direction of CASSP-funded research and training centers and the National Mental Health Association in the endorsement of less restrictive home and school-based services, family support services, and comprehensive case management services as alternatives to costly out-of-home care.[15,16]

Age and Sex Differences

Service planners generally tend to think of SED children as one class of service recipients. Our data suggest that there are significant differences in the service needs of male versus female children, and of differently aged children. These differences have serious implications for the development and implementation of systems of care for the SED population.

Our data demonstrate a substantial developmental gradient which must be considered when planning services for SED children. For school-aged children the educational setting clearly should play a large role in treatment planning. For example, school-based mental health services and supervised after school activities were distinct needs of 6–16 years olds. Therapeutic camp services and foster home placements were also highly endorsed for this subset of SED children.

For the 17–18 year old group, there is a great need for services which prepare youth for the transition for adulthood, such as training in independent living skills, vocational services, and supervised independent living arrangements. The need for outpatient and inpatient alcohol/drug treatment also peaks for this age cohort.

Sex differences in service need were less obvious than the developmental differences. Males demonstrated a greater need for transitional "adult" services, for alcohol and drug-related treatment, and for group home placement. Females, on the other hand, were in greater need of outpatient therapy, foster home placements, respite care for their families, supervised living arrangements, and protective services. The services endorsed for males focused on home removal, treatment of addiction, and preparation for entry into the work force. Services endorsed for females focused on therapy, assistance to the girls' families, and protection from further harm. More research is needed to confirm whether these gender-linked service needs are valid indices of what SED children need in their treatment, or merely reflections of sex-based stereotypes of SED children's vulnerabilities and responsiveness to various treatments and interventions.

TABLE 2—Significant Age and Gender Effects

Service	Direction of Effect
Age Main Effects	
Therapeutic preschool	1 > 2,3
Respite care	1 > 2,3
Therapeutic camp	1 > 2 > 3
Transportation to needed services	1,2 > 3
Intensive in-home services	1,2 > 3
School-based mental health services	1,2 > 3
Foster home placement	1,2 > 3
Supervised after school activities	1,2 > 3
Group home placement	2,3 > 1
Outpatient alcohol/drug treatment	3 > 2 > 1
Inpatient alcohol/drug treatment	3 > 2 > 1
Training in independent living skills	3 > 2 > 1
Vocational services	3 > 2 > 1
Supervised independent living arrangement	3 > 2 > 1
Gender Main Effects	
Outpatient therapy	F > M
Foster home placement	F > M
Respite care	F > M
Supervised independent living arrangement	F > M
Child protective services	F > M
Outpatient alcohol/drug treatment	M > F
Inpatient alcohol/drug treatment	M > F
Group home placement	M > F
Training in independent living skills	M > F
Vocational services	M > F
Supervised after school activities	M > F

Note.
Group 1 = 6–11 year olds
Group 2 = 12–16 year olds
Group 3 = 17–18 year olds
All main effects $p < .01$

Study Limitations

Limitations of this study lie in its narrow focus on professional respondents, the non-random nature of the final sample, and the survey's exclusion of preventive services.

First, our survey respondents were exclusively the professional care givers of SED children. This clearly provides only one perspective on what a given child's treatment needs may be. Family members, school personnel, and other persons in the child's support circle are equally valid respondents. Nevertheless, given that the focus of the survey was on SED children who were "system veterans" with well-known and extensive patient histories, input from the primary professional care givers was a good starting point. This in no way precludes the need to vigorously solicit the participation of parents and teachers in the development of individual children's treatment plans and larger scale planning endeavors.

Second, for the analyses of service need data, the ultimate data base of SED children was a nonrandom approximate 50 percent of our original random sample of children. Cases were eliminated for three possible reasons: 1) the survey form was not returned; 2) the survey form was returned, but was missing critical information; or 3) the child was under 6 years of age. Ideally, our final sample would have been of a completely random nature. However, because the subjects were unidentified to the investigators, and since no special services were offered to children as a result of survey responses, the final sample appeared not to be biased systematically. Case worker-related factors, such as excessive paperwork demands, rather than child-related factors, appeared to contribute to the rate of subject loss.

Third, because children who were already classifiable as SED were the subjects of this study, primary preventive services were completely excluded from the range of services offered as possible service needs on our data collection instrument. Readers must keep in mind that the services highlighted as "most needed" by our survey represent the narrow but critical range of mental health and other services for those children who are already identified as SED. Services such as community education, parent training, teen parenting classes, substance abuse prevention, advocacy, early childhood intervention, etc., are very important elements of a comprehensive strategy to combat severe childhood mental illness. ☐

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Errata

In: Robinson JC, Paxman DG. OSHA's four inconsistent carcinogen policies. *Am J Public Health.* 1991;81:775–780. Public Health and the Law.

The names of the authors, J. C. Robinson and D. G. Paxman (and not the name of the column editor), should have been listed with the article's title on the table of contents. We regret the error.

* * *

In: Mandelblatt J, Andrews H, Kerner J, Zauber A, Burnett W. Determinants of late stage diagnosis of breast and cervical cancer: the impact of age, race, social class, and hospital type. *Am J Public Health.* 1991;81:646-649.

The word "non-public" should read "public" in three different places as follows: in Table 3, Profile 5; in Table 5, Profile 5; and in Table 5, Profile 6. In Table 5, footnote a), ".67" should read ".53." In footnote b) of the same table, ".35" should read ".13."

Analysis of US Child Care Safety Regulations

Carol W. Runyan, MPH, PhD, Diana E. Gray, MPH, Jonathan B. Kotch, MD, MPH, and Matthew W. Kreuter, MPH

ABSTRACT

Background: With 1.9 million US children cared for in organized group child care, the safety of these children is a public health concern. In the absence of federal policy, each state has developed its own day care safety regulations.

Methods: After creating a set of 36 criteria from three sets of national guidelines, we assessed the safety regulations of 45 states. With a mailed survey of state day care regulatory personnel, we examined the processes of formulating and implementing safety policy in 47 states.

Results: For 24 of the 36 items, more than half the states' regulations were below the criteria or failed to mention the topic. Most notable is the inattention to playground safety, choking hazards, and firearms.

Conclusion: The uneven quality of regulations may be a reflection of a regulatory process that is fragmented, with many different groups sharing authority and with limited involvement of injury prevention specialists. (*Am J Public Health*. 1991; 81:981-985)

Introduction

According to data from the winter of 1984–85, 1.9 million United States children of working mothers were cared for in organized group day care, and an additional three million received regular care in private homes other than their own.[1]

The importance of injuries to preschoolers is well-documented,[2–4] as are the patterns of some specific injury problems including poisoning,[5,6] aspiration,[7,8] falls,[2] drowning,[9] electrical shocks,[10] scalds and burns,[11] and injuries associated with playgrounds[12] and firearms.[13,14] Only a few studies have examined injury in US day care settings.[15–21] In the most comprehensive report to date, Sacks, *et al.*[15] found that 58 percent of all day care center injuries result from falls. Falls from playground equipment account for 48 percent of serious injuries, with the playground surface an important determinant of injury severity.

Despite the large numbers of children in organized child care, the incidence of childhood injury, and several legislative efforts, no federal day care regulations exist except for centers on military bases (70,000 children) and for Head Start programs (489,000 children).[16,17] Instead, each state formulates and enforces its own regulations. State day care regulations address many issues including disease transmission, hygiene, provision of nutritious and safe foods, provision of an environment which facilitates the child's mental, physical, and emotional growth, as well as injury prevention.

Our goal was to assess the content of state safety regulations and document how they are established and enforced.

Methods

The study consisted of three elements: 1) development of the set of model criteria to assess the regulations; 2) content analysis of the regulations; and 3) a questionnaire documenting day care regulatory processes in the states.

The reference criteria were drawn from three sets of nationally recognized guidelines: Massachusetts' Department of Public Health, *Safe Daycare: A Teacher's Guide for Creating Safe Environments for Preschool Children*[10]; American Academy of Pediatrics (AAP) *Health in Daycare: A Manual for Health Professionals*[24]; National Academy of Early Childhood Programs, *Accreditation Criteria and Procedures*.[23] From a combined list of 130 injury-related criteria, 36 specific items were selected by a panel of four injury specialists including three of the authors plus one independent reviewer. Selections were based on individual judgments of the panelists about the potential importance of the injury problem for young children but with an aim of including a diverse, if limited, set of categories. We specifically focused on regulations addressing environmental features of the facilities and center procedures rather than issues such as staff hiring and training practices or staff:child

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From the University of North Carolina School of Public Health, Departments of Health Behavior and Health Education (Runyan, Gray, Kreuter) and Maternal and Child Health (Kotch), and Injury Prevention Research Center (all authors). Address reprint requests to Carol W. Runyan, MPH, PhD, Director, UNC Injury Prevention Research Center, Rosenau Hall CB 7400, Chapel Hill, NC 27599-7400. This paper, submitted to the Journal March 9, 1990, was revised and accepted for publication January 23, 1991.

ratios because the relationship between staffing patterns and injury potential is less well documented.

Content Analysis

The content analysis consisted of a comparison of each state's regulations to the 36 items. Each was rated as: 1) exceeding the criterion; 2) meeting the criterion; 3) being below the criterion; 4) not mentioning the content area; or 5) referring to other state regulations not specific to day care and not available to us (e.g. state building or fire codes). In the context of this study, we refer to day care as defined by the American Academy of Pediatrics[24] as "care, supervision, and guidance of a child or children, unaccompanied by parent or other legal custodian, on a regular basis, for periods of less than 24 hours." The content analysis, however, was restricted to analysis of regulations pertinent to day care center facilities. We excluded consideration of regulations specific to family day care as defined by the AAP; that is, "situations in which fewer than six children are cared for in the caregiver's home."[24] None of the criteria chosen delineated differences for specific age groups, although several of the criteria referred to issues specific to an age group (e.g. regulations about cribs).

Questionnaire

The questionnaire about the regulatory process and compliance was sent to the state agency responsible for regulating child day care at the same time the regulations were requested, by letter, from each state. The 15-item self-administered survey addressed the development and monitoring of safety regulations. Nonrespondents were phoned five weeks after the initial mailing and reminded of the survey.

Results

Analysis of Regulation Content

Regulations were obtained from 45 states and examined in comparison to the 36 criteria. Evaluations of the regulations, grouped by category, are presented in Table 1, demonstrating the proportion of states by whether each regulation exceeds, meets, or is below the criterion or if the topic is referred to in another code or not mentioned.

There were only five of the criteria (13.9 percent) that at least half the states met whereas 15 of the 36 items (41.7 percent) were not mentioned by half or more of the states.

For certain topics such as fire prevention, the state regulations frequently did not list specific requirements but referred to building codes or other state policies. Consequently, comparison to the criteria was impossible. For four of the six standards addressing fire and burn prevention, over half the states referred to other documents. At least 10 percent of the states referred to other regulations for swimming and wading pools, elevated areas, right hand descending stair railings, and motor vehicle passenger restraint.

Surprisingly, no state specified maximum heights for playground equipment, 96 percent neglected to mention playground surfacing, and 91 percent did not specify an eight-foot clearance between playground equipment.

There also were gaps in addressing indoor hazards. For example, 62 percent of the states were below the criterion in their restriction of objects that might cause choking, and no state specifically mentioned the most common causes of choking: hot dogs, peanuts, and grapes.[6,7] Only 26 percent of the states met the criterion for proper storage of poisonous substances, and a mere 11 percent specified that facilities not allow loaded firearms.

Although 55 percent of the states met or exceeded the criterion for tap water temperature, 38 percent failed to mention the issue. Eighty percent failed to meet the criterion that no free-standing space heaters be used.

Although 82 percent of states required that facilities have a first aid kit at the center at all times, only 18 percent required that at least one staff member be certified in cardiopulmonary resuscitation (CPR).

Subsequent to this analysis, the American Public Health Association (APHA) and the American Academy of Pediatrics (AAP) have drafted joint standards for child day care policy, including safety.[26] All of the 36 items included in our analysis are included in the AAP/APHA guidelines. For three of the 36 items, the AAP/APHA recommendations were more stringent than those used in this analysis, while for six others the AAP/APHA recommendations were slightly less stringent. For example, our analysis included a criterion that poisonous substances and medicines be kept in locked storage; the corresponding AAP/APHA guidelines indicated that there be "inaccessibility" of these substances.

The standards for Head Start programs were last revised in 1975 and indicate that if facilities meet or exceed state or local licensing requirements for fire, health, and safety this will "be accepted as prima facie compliance."[27] Our review of the 17 Head Start standards pertaining to safety indicated that none was more rigorous than the 36 criteria we considered.

Analysis of Perspectives on the Regulatory Process

Survey responses were received from 47 states (missing are Florida, Louisiana, and Maryland). Only four states *required* that injury prevention specialists be involved in developing regulations, although 16 states reported that injury prevention specialists are "routinely involved."

Safety regulation compliance by child day care centers was perceived to be good, with ratings of "better than average" reported by respondents from 26 states (55 percent). In contrast, compliance on the part of family day care homes was perceived to be "better than average" in 16 states (34 percent).

Respondents in 23 states indicated that at least one center had been closed in the prior year for safety violations and 19 reported that at least one family day care home had been closed. The types and number of inspections conducted by states varied widely. Types of inspections conducted included fire, gas pipe, sanitation, licensing, and routing monitoring; number of inspections per facility per year ranged from zero to six. Twenty-four respondents indicated that child day care centers in their states were inspected, on average, one or fewer times per year (e.g. every other year).

Those responsible for inspecting the facilities vary by state, with the jurisdiction for inspection shared by seven different types of groups or agencies. In some states, several different agencies were indicated as being responsible for development and/or enforcement of the regulations. In addition, respondents suggested that improvements were needed in: training for licensing and enforcement personnel; transportation policies; and attention to pets, electrical hazards, lead paint, and playground surfacing.

Discussion

Several limitations in the research should be noted. The content analysis was restricted to 36 specific day care center standards and to the 45 states providing

TABLE 1—Proportion of States by Level of Adherence to Criteria (n = 45)*

Criteria	Level of Adherence				
	Exceeds Criterion	Meets Criterion	Below Criterion	Topic not Mentioned	Refers to Other Code
Outdoor Hazards					
Wading pools emptied when not in use	2.2	24.4	11.1	44.5	17.8
All pools should be fenced & gates locked	6.7	20.0	17.8	24.4	31.1
Fences, natural barriers around play areas near busy streets, unsafe areas	0.0	73.3	6.7	20.0	0.0
Play equipment anchored, anchors below ground, no exposed concrete	0.0	6.6	35.6	57.8	0.0
At least 8 feet between all playground equipment	0.0	0.0	8.8	91.2	0.0
Maximum height of playground equipment 6 feet	0.0	0.0	0.0	100.0	0.0
8 inches or more of a resilient surface under playground equipment	0.0	0.0	33.3	66.7	0.0
Continuous maintenance of loose play surface materials	0.0	2.2	2.2	95.6	0.0
Flexible, light-weight, non-cutting swing seats	0.0	4.4	0.0	95.6	0.0
Structure of Facility					
Window openings limited to six inches	0.0	4.4	4.4	91.2	0.0
Elevated areas (landings, balconies, porches) provided with railings, adaptations, or approved safety gates	0.0	37.8	13.3	35.6	13.3
Right hand descending railing secured at child height	2.2	0.0	37.8	48.9	11.1
Above ground windows accessible to children must be constructed, adjusted, or adapted with window stops, screens, or grills	0.0	20.0	31.1	46.7	2.2
Indoor Hazards					
Non-metallic, tight-fitting caps in electric receptacles that are not in service in areas where young children are in care	0.0	55.5	0.0	44.5	0.0
Crib sides always up, in locked position when occupied	0.0	4.4	6.7	88.9	0.0
Toys, furnishings and equipment should be non-toxic, lead-free	0.0	48.9	33.3	17.8	0.0
Toys, furnishings and equipment should be free of rust, splinters and sharp corners	0.0	64.5	22.2	13.3	0.0
Toys, furnishings and equipment should be free from objects small enough to swallow	2.2	35.6	4.4	57.8	0.0
No peanuts, chunks of hot dogs, grapes	0.0	0.0	0.0	100.0	0.0
All potentially poisonous or hazardous substances kept in locked storage space, inaccessible to children	0.0	24.5	62.2	13.3	0.0
Rx/Non Rx medications kept in a designated locked place, inaccessible to children	0.0	28.9	60.0	8.9	2.2
No loaded gun permitted in child care facility	8.9	2.2	8.9	80.0	0.0
Fire/Burn Prevention					
Fire/Smoke Detector provided	2.2	37.8	2.2	0.0	57.8
Fire/Smoke Detector tested monthly (weekly)	2.2	6.7	2.2	17.8	71.1
Fire extinguishers provided	0.0	42.3	4.4	2.2	51.1
Fire extinguishers tested monthly (not specified)	0.0	4.4	15.6	15.6	64.4
Free standing space heaters should not be used	0.0	20.0	17.8	60.0	2.2
Hot water temperature should not exceed 120 degrees F at outlet	22.2	33.3	6.7	37.8	0.0
Emergency Procedures					
First Aid Kit in center at all times, wherever, whenever children are in care	0.0	82.2	8.9	8.9	0.0
At least one staff person on premises who is currently certified in CPR	2.2	15.6	8.9	73.3	0.0
Supervision					
Adult supervision of children at all times	0.0	91.1	8.9	0.0	0.0
No infant left unattended while bathing	0.0	11.1	0.0	88.9	0.0
Safety Education					
Safety education for parents	0.0	0.0	0.0	100.0	0.0
Safety education for children	0.0	4.4	2.2	93.4	0.0
Transportation					
All vehicle passengers seated in seats with suitable restraint system	0.0	22.2	31.1	26.7	20.0

*Data from the following jurisdictions were either not received or were not usable: Florida, Kansas, Louisiana, Maryland, Nevada and Washington, D.C.

documents. A separate analysis is needed to address the standards for the large population in family day care homes. The survey also was limited in several ways: only one representative responded from each state, and the role of these respondents may differ among states; anonymity was impossible, potentially reducing candor; and judgments of compliance were based on respondents' recall of "usual" practices which may vary widely. In addition, the scarcity of epidemiologic data about injuries occurring in day care facilities necessitates inferring, to a certain extent, that injury risks are similar to those documented for the total population of children under age five.

Nevertheless, the study does suggest the need for attention to a number of issues associated with the content, development, and monitoring of safety regulations for child care. The analysis revealed some important gaps in the regulations including inattention to playground design, firearms, and choking hazards. Given that playgrounds are the site of most reported injuries in day care centers, this oversight is especially noteworthy. Furthermore, six of the 36 standards studied were mentioned as covered under regulations other than those specific to day care by at least 20 percent of the states. It is impossible, in the context of this analysis, to assess the adequacy of those other regulations. References to other codes or standards, particularly those pertaining to fire and burn prevention, may be overlooked or misunderstood depending upon the communication and coordination among regulatory bodies or the extent to which the regulations are well-publicized. For example, knowledge about vehicle restraint laws may be more common than fire safety codes. In other instances, the failure to address an issue (e.g. firearms) may be based on the presumption that the issue is covered by other laws, a presumption that may or may not be true.

The deficiencies in the regulations reflect an inadequate regulatory process. As has been demonstrated with child automobile restraint policies, the lack of federal regulations results in diverse and inadequate state laws.[28] In the states, jurisdiction for day care safety is diffuse, with little input from injury prevention specialists. Furthermore, the relationship between regulations and the licensing process is variable among the states. Many of the regulations cannot be assessed until after a facility opens (e.g. supervision practices). Consequently, the role of inspection and relicensure may have more significance than initial licensing procedures and could benefit from further study. Interestingly, none of the national guidelines used as the basis of this analysis addressed the issue of inspection processes; however, the AAP/APHA guidelines[26] do include recommendations about inspections.

With large numbers of children in out-of-home-care other than day care centers, specific consideration of safety in family day care homes is essential. This requires attention to the delicate balance between access to affordable care on the one hand and, on the other, to the different hazards such as firearms that may be more prevalent in the home environment.

In the absence of a federal commitment, safety in child day care will remain dependent upon state actions. These efforts may be facilitated by the performance standards produced by the American Academy of Pediatrics and the American Public Health Association.[26] However helpful, they should not substitute for carefully constructed policy initiatives, preferably at the federal level, based on sound injury research. ☐

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Women Who Are Most at Risk for AIDS Perceive It Least, Say Loyola Researchers

Two Loyola University Chicago researchers have found that women who are most at risk for HIV are the least likely to perceive themselves as being at risk. They also discovered that women who are at a high risk are no more likely to be tested than those who are at a low risk.

Although other researchers had long suspected both of these facts, the Loyola team was the first to document it.

Of the 320 women surveyed, an almost equal number of minority and nonminority women indicated that they engaged in high-risk behavior. However, significantly fewer high-risk minority women agreed with statements such as "I am concerned that I will get AIDS" and "I am concerned that someone I know will get AIDS" than did high-risk nonminority women.

The research was performed last winter by Seth Kalichman, PhD, assistant professor of psychology at Loyola, and doctoral student Tricia Hunter.

"The message about the methods of AIDS transmittal clearly does not seem to be reaching high-risk minority women," said Kalichman. "Surprisingly, most of the women we surveyed had seen mass transit ads on AIDS prevention. We're hypothesizing that the ads may be lacking in cultural relevance since they did not get the desired response."

To gather information about women's perceived risk for AIDS, Kalichman and Hunter chose a unique and highly effective location. The pair approached local women of assorted ages, races, and socioeconomic classes at various public transportation sites in downtown Chicago and asked them to fill out anonymous surveys on AIDS.

"We had a cooperation rate of 93%," said Kalichman. "We didn't know what sort of reaction we would get, and the response exceeded all of our expectiations."

The Loyola researchers gave participants who completed the survey a flyer that explained the purpose of the project, gave their university telephone number for more information, and, most importantly, according to Kalichman, provided the correct answers to the survey.

Along with the debriefing sheet, the researchers handed out cards from the Chicago Board of Health which explained where the women could go to have AIDS testing done.

Kalichman and Hunter attribute a large part of their success to the fact that they approached the women during nonrush hour periods when they were simply waiting for public transportation to arrive. But more than once, women passed up a ride to finish the Loyola-funded survey—proof, the researchers think, of the women's own interest in the subject.

Because so many men asked to participate in the survey (which was restricted to women only), Loyola hopes to follow up with a second survey which will collect information about men's perceived AIDS risk.

Promoting Physical Activity and a Healthful Diet among Children: Results of a School-Based Intervention Study

Bruce G. Simons-Morton, EdD, MPH, Guy S. Parcel, PhD, Tom Baranowski, PhD, Ronald Forthofer, PhD, Nancy M. O'Hara, PhD

ABSTRACT

Background: National health objectives call for improved diet and more regular physical activity among children. We tested the effects of a school-based program to improve students' diet and physical activity behavior at school.

Methods: Two of the four elementary schools in one Texas school district were assigned to intervention and two to control conditions. The three intervention components were classroom health education, vigorous physical education, and lower fat, lower sodium school lunches. Nutrients from school lunches and the total day and the amount of physical activity students obtained during physical education were assessed as outcome.

Results: Analysis of school lunches showed declines from base line to posttest in the two intervention schools of 15.5% and 10.4% for total fat, 31.7% and 18.8% for saturated fat, and 40.2% and 53.6% for sodium; posttest values were lower in the intervention schools. Observation of physical activity during physical education classes indicated an increase in the intervention schools from baseline to posttest in the percent of time children engaged in moderate-to-vigorous physical activity from less than 10% of class time at baseline to about 40% of class time at posttest; posttest values were higher in the intervention schools than in the control schools.

Conclusions: This efficacy study demonstrates the feasibility of substantially modifying school lunches and school physical education to improve children's diet and physical activity behavior at school. (*Am J Public Health.* 1991;81:986–991)

Introduction

Childhood diet and physical activity may lead to increased risk for cardiovascular disease (CVD) in adulthood.[1] Promoting a healthful diet and regular vigorous physical activity among children is a national public health objective.[2] School is an ideal setting in which to promote healthful diet and physical activity because these topics usually are part of the health education curriculum, and federally supported food services and state-mandated physical education (PE) are widespread.[3] School food services provide nutritious meals for 73.4% of all US public school students,[4,5] but the high fat and sodium content is a concern.[6] An estimated 97% of first- through fourth-grade children in US public schools are enrolled in school PE programs for an average of 100 minutes per week.[7] However, the frequency and duration of PE classes and the amount of moderate to vigorous physical activity (MVPA) children obtain during PE may be less than recommended.[8,9]

This study was designed to influence the school environment in terms of affect on student diet and physical activity at school. The Go For Health (GFH) program included a behaviorally-based health education curriculum, fitness-oriented PE, and lower fat and sodium school lunches. We report the effects of the program on children's dietary intake and physical activity during school.

Methods

Design and Study Population

The Texas City Independent School District (TCISD) was selected for the study because it is conveniently close to the university, it has an ethnic mix of students, and it has curriculum, school lunch, and PE components. The district participated in the US Department of Agriculture school lunch program, and each school had its own on-premises kitchen, cafeteria, and food preparation staff. Third and fourth graders had PE daily, taught by PE specialists. Facilities at each school include a modern gymnasium and large outdoor playground area. The population of students was 62.3% Anglo-American, 20.9% Mexican-American, and 14.8% Afro-American. The ethnic distributions were comparable for both treatment and control conditions.

Two of the four elementary schools (kindergarten through fourth grade) were assigned to intervention and two to control conditions. Third and fourth graders were assessed annually during three spring data collection periods.

Intervention

The three GFH intervention components, based on social cognitive theory,[10] were the Go For Health Curriculum, Children's Active Physical Education (CAPE), and the New School Lunch (NSL).[11] Following the model of Charter

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Bruce G. Simons-Morton, Guy S. Parcel, Ronald Forthofer, and Nancy M. O'Hara are with the School of Public Health of the University of Texas Health Science Center, Houston, TX. Tom Baranowski is with the Georgia Prevention Institute, Medical College of Georgia, Augusta, Ga.

Requests for reprints should be sent to Bruce G. Simons-Morton, EdD, MPH, Center for Health Promotion Research and Development, School of Public Health, University of Texas Health Science Center, PO Box 20186, Houston, TX 77225.

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TABLE 1—Average Nutrients per Lunch (N = 12 Meals per School) at Baseline, Midtest, and Posttest, with 95% Confidence Intervals: Go for Health Recipe Analysis

Treatment condition	Energy (kcal)	Protein (g)	Carbohydrate (g)	Fat (g)	Cholesterol (mg)	Vitamin A (IU)	Calcium (mg)
Control school 1							
Baseline[a]							
Mean	1049.2	38.3	115.1	48.4	123.8	2722.7	535.5
Lower 95% CI	1014.3	37.4	112.7	46.0	116.5	2291.0	516.4
Upper 95% CI	1084.1	39.2	117.3	50.9	131.1	3154.4	554.6
Midtest							
Mean	997.5	36.2	107.7	46.9	113.2	2985.3	537.2
Lower 95% CI	949.3	35.2	104.5	43.4	104.5	2301.4	510.5
Upper 95% CI	1045.8	37.3	111.0	50.3	121.8	3669.2	564.0
Posttest							
Mean	898.2	35.3	100.4	39.5	94.6	2071.5	480.3
Lower 95% CI	870.4	34.4	98.2	37.8	90.4	1738.3	461.4
Upper 95% CI	926.0	36.2	102.4	41.3	98.7	2404.7	499.3
Control school 2							
Baseline[a]							
Mean	1049.2	38.3	115.1	48.4	123.8	2722.7	535.5
Lower 95% CI	1014.3	37.4	112.7	46.0	116.5	2291.0	516.4
Upper 95% CI	1084.1	39.2	117.1	50.9	131.1	3154.4	554.6
Midtest							
Mean	1038.3	37.1	114.9	47.8	112.2	2418.6	539.9
Lower 95% CI	996.4	35.7	112.4	44.9	104.2	2019.6	507.9
Upper 95% CI	1080.3	38.4	117.8	50.6	120.2	2817.7	572.0
Posttest							
Mean	943.5	37.1	103.6	42.3	111.7	2106.0	511.7
Lower 95% CI	918.6	36.1	102.2	40.6	103.4	1756.4	488.7
Upper 95% CI	968.4	38.0	105.3	43.9	119.9	2455.5	534.7
Intervention school 3							
Baseline							
Mean	1077.5	37.3	117.8	50.8	137.7	2809.4	544.0
Lower 95% CI	1037.8	36.3	114.9	48.1	129.8	2365.6	522.3
Upper 95% CI	1117.1	38.3	120.6	53.5	145.6	3253.1	565.6
Midtest							
Mean	890.4	37.7	105.3	35.4	98.1	3559.6	514.5
Lower 95% CI	856.3	36.4	102.5	33.4	91.6	3020.8	496.2
Upper 95% CI	924.5	38.9	108.1	37.4	104.7	4098.4	532.7
Posttest							
Mean	849.3	35.0	100.8	34.0	79.0	3393.0	474.7
Lower 95% CI	816.8	34.4	97.8	32.0	76.8	2916.7	462.1
Upper 95% CI	881.8	35.7	104.0	35.9	81.1	3869.3	487.2
Intervention school 4							
Baseline							
Mean	1021.0	39.3	112.2	46.1	110.0	2636.0	527.0
Lower 95% CI	856.6	37.9	110.4	43.6	101.8	2210.9	508.3
Upper 95% CI	985.4	40.6	114.4	48.5	118.1	3061.0	545.7
Midtest							
Mean	887.3	33.8	108.2	35.5	83.5	3430.8	501.5
Lower 95% CI	856.6	32.7	106.1	33.5	78.2	2903.0	481.3
Upper 95% CI	918.0	34.8	110.3	37.5	88.7	3958.6	521.7
Posttest							
Mean	840.9	35.7	96.9	34.5	79.3	3907.6	469.5
Lower 95% CI	800.7	35.0	93.0	32.1	77.1	3305.2	452.4
Upper 95% CI	881.1	36.4	100.6	37.0	81.4	4509.9	486.6

[a] Baseline for control schools imputed from the average of the intervention schools.

and Jones,[12] we employed a four-step approach to program implementation and institutionalization.[13] To foster quality implementation, the GFH staff provided ample training, consultation, and technical support to both teachers and food service staff.

The purposes of the classroom health education, which consisted of six, behaviorally based modules, were to teach knowledge and skills essential to lifelong performance of the target diet and physical activity behaviors and foster the transfer of learning about these behaviors from school to out-of-school environments.[10]

The NSL was designed to provide lower-fat, lower-sodium lunches within the context of the existing school lunch program. Food purchasing, menus, recipes, and food preparation practices were modified.

CAPE, which consisted of five, 6- to 8-week units, was designed to encourage enjoyable MVPA among children during PE classes. Each unit included two or three main cardiovascular fitness activities, such as dancing, running, aerobic games, jump rope, and obstacle courses. Each class session consisted of warmup, fitness development, cool down, and skill development or game activities.

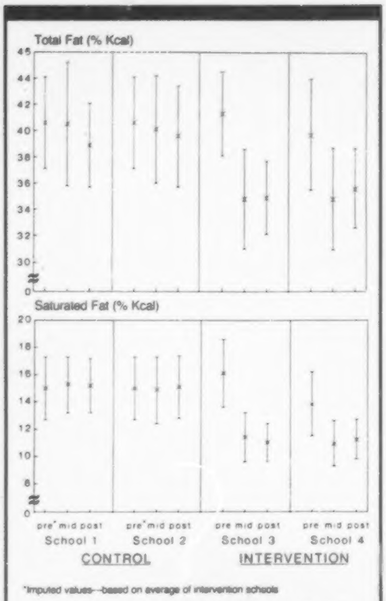

FIGURE 1—Means and 95% Confidence Intervals for Total Fat and Saturated Fat as a Percent of Calories in School Lunches (N = 12 meals per school) in Control and Intervention Schools at Baseline (Pre), after 1 Year (Mid), and after 2 Years (Post).

Measures and Procedures

Program effects were assessed by analyzing the nutrient contents of lunches, conducting 24-hour dietary recalls, and observing students' physical activity during PE.

Nutrient Analysis of School Lunches. The identical menu was repeated in March and April each year of the study. At baseline, 12 meals were selected randomly from the March and April menus and analyzed for nutrient content. Recipe analyses, based on detailed interviews with each cook in the intervention schools at baseline and in all four schools at midtest and posttest, were conducted by trained staff nutritionists and analyzed by the Nutrition Coding Center (NCC).[14]

Diet Recall. The 24-hour dietary recall interview was administered at posttest to random, cross-sectional subsamples of children, stratified by study school. Interviewers were trained according to NCC procedures and followed the NCC protocol for diet interviews. Each child was interviewed at home in the company of one parent, who was asked about ingredients, food preparation, and other items as needed.

Physical Activity. The time (in minutes) of MVPA performed by children during physical education was assessed by trained observers using the Children's Physical Activity Observation Form (CPAOF).[15] Study schools were observed on randomly selected days over a 2-month period each spring; all third- and fourth-grade classes were observed. Observers recorded on a minute-by-minute basis the type (e.g., basketball, jump rope) and intensity of physical activity. The CPAOF intensity categories have been validated against heart rate,[15] and the average heart rate for MVPA was 172.2 (SD = 21.2) beats per minute. Interobserver agreement was 97.0% (n = 57 paired observations).

Analysis

For energy and selected nutrients of potential concern in this age group, we calculated the means and 95% confidence intervals (CIs) of the repeated samples of school lunches in each school. For each separate treatment condition, we calculated the mean nutrient intakes from the 24-hour dietary interviews for selected nutrients and examined the contribution of school and bag lunches to total daily nutrient intakes.

For each school at each reporting period, we calculated the MVPA means, standard deviations, medians, and 95% CIs for the medians.[16] The median is the most appropriate statistic for purposes of comparing treatment effects because increases in the amount of MVPA on the part of only a few students could affect the mean value, obscuring true program effects. Percent of class time is preferable to minutes as an outcome because the scheduled length of PE classes varied somewhat from year to year, and the effective class time available for PE teachers and students, although similar in each school for a given year, varied somewhat from school to school due to local scheduling. The annual surveys are presented separately for each grade because most of one year's third graders became the next year's fourth graders; hence the data for each grade are not independent across years.

Results

Diet

Recipe Analysis. Table 1 includes the means and the 95% CIs for total energy (in kilocalories [1 kcal = 4.2 kJ]) and six nutrients in the 12 meals served in each study school each year. Baseline values for the control schools, which were not collected directly, were estimated by averaging the values of the two intervention schools. Declines from baseline to posttest in each school for protein, carbohydrates, total fat, cholesterol, and calcium are the product of declines in energy. The decline in calories served in the intervention schools appears to have been due mainly to a reduction in the amount of fat and carbohydrates. Despite the declines in energy, vitamin A and calcium did not decline, and posttest values for all schools exceeded the USDA school lunch requirements for each nutrient. For example, at every reporting the average intake of energy in each school exceeded 33% of the total daily requirement for energy, which is 600 to 700 kcal per meal.[17]

Shown in Figure 1 for each school is the average total fat and saturated fat as a percent of kcal served in the sample of 12 lunches. For both intervention schools the upper CIs of the posttest means were less than the pretest means, and in school 3 the upper and lower CI values did not overlap. The mean percent fat declined in school 3 by 15.5%, from 41.3% to 35.6% of calories, and in school 4 by 10.3%, from 39.7% to 35.6% of calories. Average percent saturated fat declined in school 3 by 31.7%, from 16.1% to 11.0% of calories, and in school 4 by 18.8%, from 13.8% to 11.0% of calories. The means for percent total fat and saturated fat at posttest in the control school were within the imputed pretest 95% CIs.

Figure 2 shows that the sodium content in the school food declined from baseline to posttest in school 3 by 53.6% from 2141.1 mg to 993.0 mg, and in school 4 by 40.2%, from 1730.0 mg to 1034.5 mg. These differences remained after controlling for kcals served. Sodium in the control schools did not decline from the imputed baseline value.

Dietary Recall. Table 2 shows the analysis of the posttest 24-hour recalls for the nutrients of interest. Students in the intervention group, compared with students in the control group, reported fewer calories (2094 vs 2135 [1 calorie = 4.2 J]) and less total fat (77 g vs 90 g), saturated fat (29 g vs 34 g), and sodium (2873 mg vs 3378 mg) for the total day. In comparison with control group students, intervention group students reported consuming less total fat, saturated fat, and sodium in both tray and bag lunches.

Physical Activity

Table 3 shows medians, means, and standard deviations for minutes of MVPA for observed students in grades three and four at each study school at baseline, midtest, and posttest. At baseline, mean MVPA for third and fourth graders was less than 3 minutes at each school; medians were 2.0 minutes or lower. At posttest, mean minutes had increased in school 3 to 11.7 (third grade) and 15.0 (fourth grade) and in school 4 to 16.2 (third grade) and 16.1 (fourth grade) per class.

Median minutes as a percent of class time and 95% CIs around the medians are shown for third and fourth grade in Figure 3. For both grades, the median percent of class time in MVPA at baseline was zero in three schools and less than 10% in school 3. At posttest, the control school medians were still no greater than 5% of class time. In the intervention schools, however, the posttest medians were substantially higher, 28% of class time for third graders in school 3, and 39% of class time or greater for the other intervention groups. Except for fourth grade in school 3, the upper and lower 95% CI values do not overlap with the baseline means.

Discussion

The results show strong positive program effects on children's diet and physical activity at school. At posttest, school lunches in the intervention schools provided less energy, fat, and sodium compared with baseline and with the control schools, while adequate levels of other essential nutrients were maintained. Similarly, children's physical activity in the intervention schools was much higher at posttest than at baseline and in comparison with the control schools.

Several studies have demonstrated that sodium reductions can be accomplished by intervention directed at school meals.[18,19] However, no studies have reported systematic reductions in fat and sodium in school lunches. In the GFH intervention schools, sodium content was reduced to target levels, saturated fat reductions of 31.7% and 18.8% occurred, and posttest values approached the target goal of 10% of kcals served. Total fat reductions were modest, 15.5% and 10.3%, and the target goal of 30% of kcals was not reached, indicating that it is easier, even in institutional settings, to substitute fat type than to reduce the amount of fat.

The 24-hour dietary recall data indicate that the changes in nutrients served in

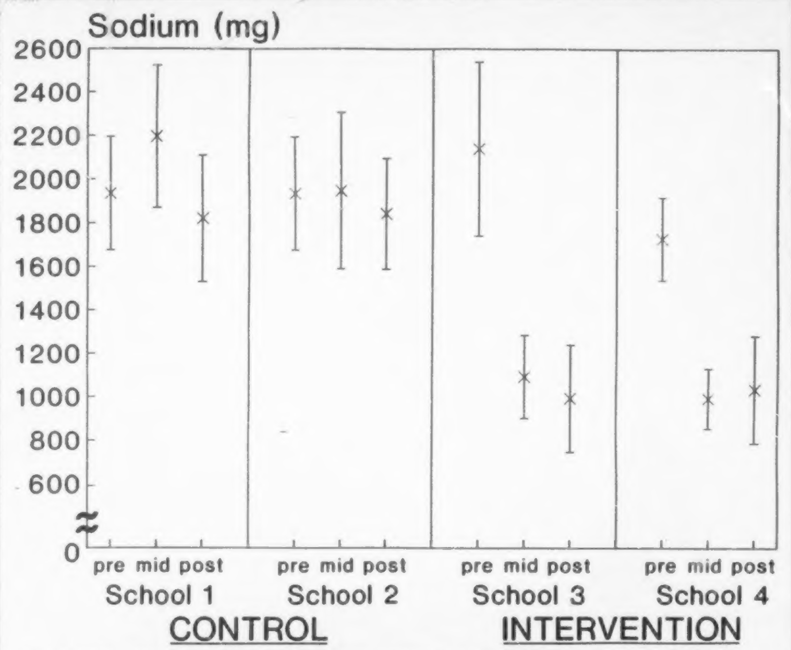

FIGURE 2—Means and 95% Confidence Intervals for Sodium in School Lunches (N = 12 Meals per School) in Control and Intervention Schools at Baseline (Pre), after 1 Year (Mid), and after 2 Years (Post).

TABLE 2—Means and 95% Confidence Intervals of Selected Nutrients from Bag or Tray Lunch Based on 24-Hour Diet Recalls in the Intervention and Control Schools

Variables	Energy (kcal)	Total fat (g)	Saturated fat (g)	Sodium (mg)
		Control		
Total day (n = 68)				
Mean	2135.3	90.0	33.9	3377.6
Lower 95% CI	1978.3	80.8	30.3	3044.1
Upper 95% CI	2292.3	99.1	37.6	3711.0
Tray lunch (n = 49)				
Mean	806.9	34.8	13.8	1283.4
Lower 95% CI	730.7	30.8	11.9	1137.2
Upper 95% CI	883.0	38.8	15.8	1429.6
Total day average (%)	37.8	40.1	40.7	39.7
SE	1.8	2.4	2.4	2.3
Bag lunch (n = 19)				
Mean	681.4	27.3	9.0	966.0
Lower 95% CI	568.8	21.5	6.9	794.8
Upper 95% CI	794.1	33.0	11.0	1137.3
Total day average (%)	35.7	34.4	32.5	33.4
SE	2.6	2.7	3.4	2.7
		Intervention		
Total day (n = 67)				
Mean	2094.5	78.6	28.5	2872.6
Lower 95% CI	1922.6	70.0	25.0	2554.2
Upper 95% CI	2266.5	87.2	32.0	3190.7
Tray lunch (n = 48)				
Mean	790.2	29.7	10.0	909.1
Lower 95% CI	676.9	25.6	8.3	756.6
Upper 95% CI	903.4	33.8	11.6	1062.7
Total day average (%)	36.5	38.0	35.4	32.5
SE	1.9	2.2	2.3	2.0
Bag lunch (n = 19)				
Mean	657.7	25.1	7.8	789.5
Lower 95% CI	516.8	18.8	5.6	645.5
Upper 95% CI	798.6	31.3	10.0	933.5
Total day average (%)	36.1	37.2	32.6	31.3
SE	2.7	2.8	3.2	2.4

TABLE 3—Average Time (in Minutes) of MVPA Observed during Physical Education in Control and Intervention Schools

Treatment Condition	Third Grade				Fourth Grade			
	N	Median	Mean	SD	N	Median	Mean	SD
				Control				
School 1								
Baseline	32	0.0	0.4	0.9	41	0.0	0.3	0.7
Midpoint	27	1.0	2.8	5.0	36	0.5	1.1	1.5
Posttest	37	2.0	2.9	3.5	35	1.0	1.5	1.9
School 2								
Baseline	34	0.0	0.5	1.1	26	0.0	0.4	0.9
Midpoint	24	1.0	2.0	2.2	10	0.0	2.2	3.4
Posttest	36	2.0	3.5	3.9	51	1.0	4.5	5.9
				Intervention				
School 3								
Baseline	31	2.0	2.0	1.4	45	2.0	2.4	1.9
Midpoint	34	1.5	2.6	2.6	24	2.5	6.8	7.4
Posttest	38	10.5	11.7	7.5	26	20.0	15.0	10.6
School 4								
Baseline	32	0.0	2.8	5.0	36	0.0	1.9	3.6
Midpoint	38	4.5	4.9	4.3	24	7.0	8.0	4.4
Posttest	58	16.0	16.2	7.8	49	16.0	16.1	7.4

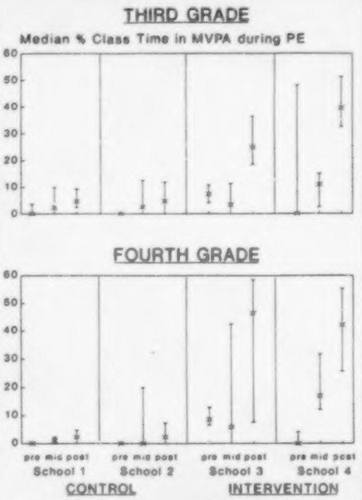

FIGURE 3—Medians and 95% Confidence Intervals for Percentage of Physical Education Class Time in Moderate to Vigorous Physical Activity (MVPA) in Control and Intervention Schools at Baseline (Pre), after 1 Year (Mid), and after 2 Years (Post); Third and Fourth Grade.

school lunch contributed to improvements in total daily dietary intake of calories, total fat, saturated fat, and sodium. Children did not make up in other meals for reductions of fat and sodium in school lunches. Energy, calcium, and vitamin A remained adequate at posttest in all schools, suggesting that reductions in fat can be accomplished without undue reductions in other essential nutrients.

Several studies have demonstrated that it is possible to modify PE to improve cardiorespiratory fitness.[20-22] Our results indicate that it is possible to increase time devoted to MVPA from less than 10% to more than 40% of available class time. At posttest, the typical student in the intervention schools obtained about 16 minutes of MVPA per daily PE period, or 80 minutes of MVPA per week. It must be noted that these modest results were obtained after 2 years of intervention that involved one staff person virtually full time in the provision of intensive training of PE teachers. Because PE accounts for nearly half of children's total daily physical activity,[23] the extremely low level of MVPA observed in all schools at baseline is a cause for concern, particularly if future research determines that low levels of activity during PE are prevalent in American schools.

This is the first US study to demonstrate the potential of altering the school environment to promote a more healthful diet and more vigorous physical activity among children, although at least one successful European study has been reported.[24] Strengths of the present study include planned implementation of theory-based individual and organizational interventions, two intervention and two control schools within the same district, and objective measures of diet and physical activity. Limitations of the study include the lack of baseline recipe data for the control schools, nonrandom assignment of schools to treatment conditions, and the small number of study schools.

The results of this efficacy study suggest that school is a potentially important site for improving children's diet and physical activity and that such changes are possible. In our experience, however, schools do not change readily, and substantial staff training should accompany policy changes regarding healthful school lunch and vigorous physical education. Before widespread implementation of these changes occurs, effectiveness studies with larger numbers of schools are needed to determine the ability to generalize these approaches and to assess cost effectiveness. ☐

Acknowledgments

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National Primary Care Nurse Practitioner Symposium Issuing Call for Abstracts

The 17th National Primary Care Nurse Practitioner Symposium is accepting abstracts of research, innovative clinical practice/models and roles, and other professional papers relating to primary care nursing. Abstracts are being accepted for presentation (plenary) or poster sessions. This symposium, sponsored by the University of Colorado Health Sciences Center School of Nursing, will be held July/early August, 1992, in a Colorado mountain resort.

Please send three typed copies (300 words or fewer), limited to one page, single spaced, with 1-1/4 inch margins. Abstracts must be photocopy-ready and include title. A separate page with author contact information may be used. Format includes title, brief statement of purpose, methods, conclusions, and clinical application(s) or, if not a research abstract, brief summary of project/paper. Include author vita and indicate interest in presentation (plenary) and/or poster session.

Preference will be given to abstracts with strong clinical applicability, innovative role development models and papers with future-oriented perspective.

Presenters are provided tuition waiver for registration the day of their presentation, but are responsible for their own travel and expenses. Deadline is October 14, 1991.

Contact Ellen Lemberg, RN,C, MS, Nurse Practitioner Symposium Office, School of Nursing, University of Colorado Health Sciences Center, 4200 East 9th Avenue, Box C287, Denver, Colorado 80262. Tel: 303/270-7436.

Attitudes toward Abortion among Parents of Children with Cystic Fibrosis

Dorothy C. Wertz, PhD, Janet M. Rosenfield, MS, Sally R. Janes, MSW, and Richard W. Erbe, MD

ABSTRACT

Background: DNA prenatal diagnosis for cyctic fibrosis (CF) has been available for parents of affected children since late 1985.

Methods: Using anonymous questionnaires, we surveyed 395 parents of children with CF at 12 New England CF centers with regard to 12 maternal or family situations and 11 fetal characteristics; 271 (68%) responded.

Results: The majority supported legal abortion in the first trimester for all 23 situations; 58% would abort for severe mental retardation (MR), 40% would abort for a genetic disorder leading to death before age five years, 41% for a child bedridden for life, 35% for moderate MR, 20% for CF and 17% for a severe incurable disorder starting at age 40 years. Few would abort for a disorder starting at age 60 years, for genetic susceptibility to alcoholism or for sex selection. Variables most strongly related to abortion for CF were attitudes of spouse, respondent's siblings, and CF doctor toward abortion for CF as well as infrequent attendance at religious services.

Conclusions: Prenatal diagnosis may not reduce substantially the number of CF births to parents of CF children because most do not accept abortion for CF. (*Am J Public Health.* 1991;81:992–996)

Introduction

Parents of children with genetic disorders are frequently ambivalent toward selective abortion of affected fetuses and may regard the abortion as a rejection of their affected child. Although some studies report that the majority of these parents do not consider abortion acceptable for the disorder in question,[1–8] other studies indicate that such parents may be more willing to abort than the general population.[9–10]

Attitudes toward abortion follow a spectrum of perceived severity, severe mental retardation (MR) being the most crucial factor. Most women (94% to 97%) with prenatally diagnosed fetal abnormalities involving severe MR (e.g., trisomies 13, 18, or 21) choose abortion.[11] Fewer women (62%) choose abortion for sex chromosome abnormalities (e.g., 45, X Turner syndrome or 47, XXY Klinefelter syndrome) that are not ordinarily associated with retardation.[12] Severe physical disability ranks next. Among women undergoing MSAFP (maternal serum α-fetoprotein) screening for neural tube defects, 57% would abort if there "was a very good chance that the child would be born with a serious physical handicap [sic], for example, blindness or paralyzed legs."[13]

Some common genetic disorders such as cystic fibrosis (CF) involve neither MR nor significant physical limitations, but carry a reduced life expectancy and increased morbidity. CF is the most common autosomal recessive disorder affecting Caucasian populations, with an incidence of 1 in 2000 to 3000 births. About 30 000 individuals in the United States are affected. Improved palliative treatment has extended life expectancy and improved quality of life. In 1988, 75% survived into their late teens, half reached the age of 26, and about 40% reached the age of 30, though few survived to their fifth decade.[14]

Since late 1985, identification of deoxyribonucleic acid (DNA) probes linked to the CF gene has allowed accurate prenatal diagnosis for families who have had a previous child with CF and for whom specimens for the affected child and both parents can be analyzed in parallel with amniotic fluid or chorionic villi (CVS) specimens.[15–16] In other nations the majority of such families (65% in Belgium,[17] 52% in Wales[18]) would abort a CF fetus. In France, parents chose to terminate 53 of 56 CF fetuses following prenatal diagnosis.[19] There are no comparable studies in the United States. Our specific aims were (1) to find out how parents of affected children viewed abortion for CF, and (2) to see where parents placed CF in a wider spectrum of situations that might justify abortion.

Methods

Questionnaire Content

Our questions on abortion were part of a larger study of psychosocial factors

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Dorothy C. Wertz is with the Health Services Section, School of Public Health, Boston University. Janet M. Rosenfield is with the Shriver Center for Mental Retardation, Waltham, Mass. Sally R. Janes and Richard W. Erbe are with the Division of Human Genetics and Department of Obstetrics/Gynecology, Children's Hospital, Buffalo, NY.

Requests for reprints should be sent to Dorothy C. Wertz, PhD, Health Services Section, School of Public Health, Boston University, 80 East Concord Street, Boston, MA 02118.

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affecting use of prenatal diagnosis for CF in six New England states. We anticipated that parents might be less likely to abort for CF than for severe MR, severe physical disability, conditions that threatened the mother's life or health, and rape or incest, and more willing to abort for CF than for late-onset or cosmetic conditions.

We asked about attitudes toward abortion in 23 situations, including 12 maternal or family situations and 11 conditions affecting the child. Respondents were asked to give one of three responses: (1) "I would have an abortion," (2) "I would not have an abortion, but it should not be prohibited for others," or (3) "I think abortion should be prohibited by law." We included situations described in opinion polls (mother's life, rape, incest), or used as rationales for abortion (maternal age under 15, mother's career, financial burden, family completed). In describing fetal characteristics, we avoided listing names of specific disorders and instead briefly described the child's condition (e.g., instead of trisomy 18, "severe mental retardation: child unable to speak or understand"; instead of Tay-Sachs, "severe genetic disorder leading to death before age 5"; instead of Huntington disease, "severe painful disorder starting at age 40, incurable"). We included several conditions that are not presently diagnosable prenatally, such as susceptibility to alcoholism, "severe incurable disorder at age 60" (Alzheimer disease), and "severe, untreatable obesity." We included obesity because of concerns that, if ever given the opportunity, some parents may make prenatal selections on cosmetic grounds, such as hair, eye, or skin color, or stature.[20] Finally, we included sex selection because, despite much publicity and two recent surveys of physicians,[21-23] there are no surveys of parental attitudes.

Previous studies have found associations between abortion attitudes and religion,[9,10,24] religiosity,[13,25] education,[10,25,26] income,[10] previous elective abortion,[9] and interpretation of risk,[9] but a recent study of CF families in the United Kingdom found no association between abortion attitudes and social class.[18] The 15-page questionnaire included the above variables and sections on reproductive history, reproductive plans, health of the affected child, expectations about the child's activities 5 and 10 years from now, knowledge about new genetic tests, and standard sociodemographic data. Most questions called for categorical (yes-no) responses. Five-point Likert scales were used to describe the interpretation of risk and the attitudes of spouses, members of extended family, and medical professionals toward abortion for CF, and (in a separate question) support of those persons carrying a CF fetus to term.

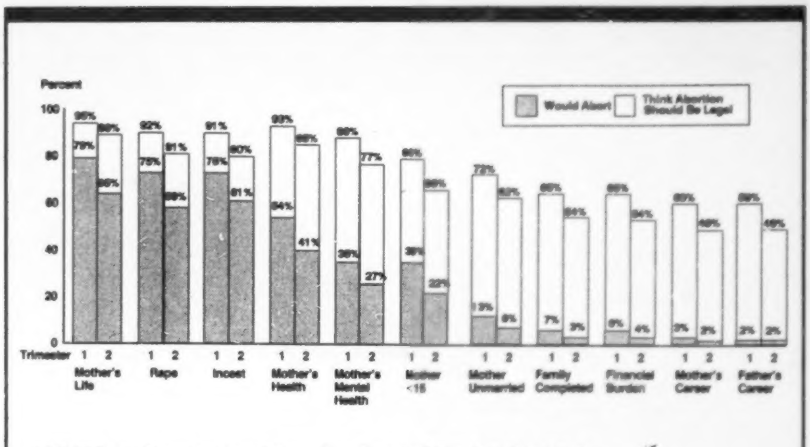


FIGURE 1—Abortion for Maternal or Family Situation (N = 271).

There are liabilities in studying attitudes rather than actual behaviors. Low prevalence and the practical difficulties of obtaining abortion histories limited us to a study of attitudes. Some studies have found attitudes strongly related to subsequent behavior.[24,27]

Distribution

Starting in January 1989, all 12 CF centers in New England distributed the questionnaires to parents of children enrolled as patients. Depending on the clinic, questionnaires were either given out at regular clinic visits over a 4-month period (most CF patients visit every 3 months) to be returned by mail, or were mailed directly to families. The two modes of administration produced almost identical response rates. In order to ensure anonymity, all questionnaires were returned directly to project staff at the Shriver Center, not to the clinics.

Data Analysis

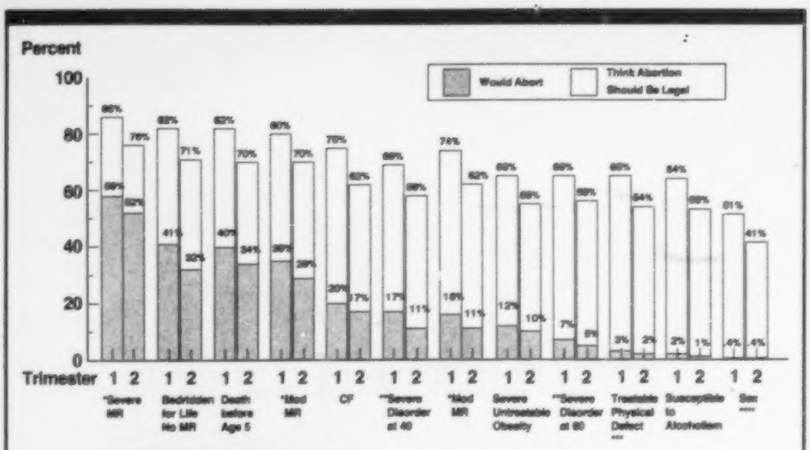
Questionnaire data were entered into Statistical Package for the Social Sciences (SPSSX) and Statistical Analysis System (SAS) programs. Using the statement "I would have an abortion" in the first trimester as the dependent variable, we looked for associations with sociodemographic variables, child's health, future expectations, interpretation of risk, and extended family's attitudes toward abortion. The five-point Likert scales were reduced to three points in order to eliminate small cells for "strongly approve." Analyses of subsets by gender, religion, and religiosity (weekly attendance) were conducted to rule out the possibility of interaction.

For each of the 23 abortion questions, all variables that associated, at the zero-order level, with "I would have an abortion" were entered into a stepwise logistic regression. This method orders each independent variable in terms of its strength of association with the dependent variable, while controlling for other variables. For each independent variable that was significantly associated, the method provided an odds ratio for personal acceptance of abortion.

Profile of Respondents

Of 395 parents asked to participate, 271 (68%, 228 families) responded. Their children's health status corresponded generally with the national clinical profile of CF.[11,28-29] The children's median age was seven; 90% had pancreatic involvement; 73% required chest physical therapy, including 46% who required it from one to five times daily; and 42% had been hospitalized within the last year.

Parents appeared typical of CF families in other studies: median age was 35; 36% were college graduates and median income was $30 000 to $40 000. These figures were comparable to an earlier California study of CF families' attitudes toward prenatal diagnosis.[30] Seventy-eight percent were living with the affected child's other parent, and 57% were Catholic, figures comparable to a New York City study.[31] The high proportion of Catholics may result from greater frequencies of the CF gene in the Irish, Italian, and French populations found in New England.[14,31] Of respondents, 73% were fe-



*Severe MR: Unable to speak or understand
Moderate MR: Could communicate, not live independently
Mild MR: Could hold job and live independently
**Develop severe painful disorder, incurable
***Cleft palate, nearsightedness
****Child is not sex desired by parents

FIGURE 2—Abortion for Fetal Characteristics (N = 271).

male and 13% had had a previous elective abortion. In all, 52% had been surgically sterilized (compared with 62% in a previous study[30]), 15% were either beyond reproductive age or no longer living with the child's other parent, and 33% were still fertile and at risk for having another child with CF.

Although 27 persons did not respond to the question about abortion for CF in the first trimester or to other questions about abortion, there were no statistically significant differences between respondents and nonrespondents.

Results

Responses are described in Figure 1 (maternal or family situations) and Figure 2 (fetal characteristics). The majority believed that abortion should be legal in all 23 situations during the first trimester and in 20 of the 23 situations during the second trimester (white bars). For all situations, fewer respondents approved legal abortion in the second trimester than in the first trimester, with an average difference of 11 percentage points.

There was a steady gradation in personal acceptance of abortion (shaded bars), running from "mother's life in danger" down to "child would interfere with father's career" (Figure 1) and from "severe mental retardation" to "child is not of sex desired by parents" (Figure 2). Severe MR was the only fetal characteristic for which the majority (58%) would themselves abort (Figure 2). For CF itself, 20% would abort in the first trimester, and 17% in the second trimester.

In both Figures 1 and 2, 80% to 100% of those represented in a shaded bar are also represented in the shaded bars to the left. For example, of those who would abort for CF, 86% would also abort for moderate MR, 88% if the child would be bedridden for life, and 100% if the child would die before age five or have severe MR. An average of 9% fewer women would abort in the second trimester than in the first trimester for maternal or family situations, and an average of 5% fewer women would abort in the second trimester for fetal characteristics. Men made less differentiation between trimesters (6% for maternal situations and 1% for fetal characteristics).

Characteristics associated at the zero-order level with personal willingness to abort are presented in Table 1. Men were more likely than women to say either that they would abort or that abortion should be prohibited by law. Men eschewed the middle ground of not aborting but permitting it for others. Gender was not related to any responses concerning fetal characteristics.

For all maternal or family situations and all fetal characteristics except sex selection, attitudes toward abortion were associated with respondent's perception of attitudes of spouse, mother, father, and siblings toward abortion of a fetus with CF. The majority thought that their spouses (65%), mothers (60%), fathers (55%), and siblings (51%) would disapprove. These views were closely interrelated, with Pearson correlation coefficients of >.6. Family attitudes toward abortion for other disorders were not surveyed.

In addition, the following variables (not reported in Table 1) were associated with abortion for CF: perceived approval of abortion for CF by father-in-law or mother-in-law, approval of abortion for CF by affected child or by other children, approval of abortion for CF by CF doctor or genetic counselor, and perceived absence of support from spouse, family doctor, or CF doctor for carrying a fetus with CF to term.

Not associated with any attitudes toward abortion, including abortion for CF, were age, fertility status, marital status, number of children, CF child's health, future expectations about the CF child, cost of the child's illness, knowledge about prenatal diagnosis, interpretation of genetic risk, and reproductive intentions. No sociodemographic variables were significantly related to abortion for obesity, severe disorder at 60, susceptibility to alcoholism, treatable defect, or sex selection.

Stepwise logistic regression analyses were performed for all abortion situations in Table 1. Characteristics strongly and independently related to willingness to abort ($P < .10$) are indicated.

Some general patterns emerged from the regression analyses. Religious background was related largely to maternal or family situations, where traditional Catholic social values may remain, rather than to abortion for fetal characteristics. In many situations, those with higher education were more willing to abort than those with less education.

Family attitudes were among the strongest factors related to personal abortion decisions. For persons living in a family context, it is difficult to make an abortion decision that contradicts the values of their spouses and extended families. Respondents' perceptions of their mothers' views were especially important for maternal or family situations. For spousal views, respondents' perceptions were accurate in the 43 families in which both spouses responded.

Regression results for abortion for CF and severe MR are reported in Tables 2 and 3. Table 2 shows that parents' perceptions of the views of their CF doctors toward abortion for CF were significant. In all, 15% thought their CF doctors would approve, 45% thought they would disapprove, and 40% thought they would be neutral.

TABLE 1—Psychosocial Characteristics Associated with Personal Willingness to Abort in Selected Situations (N = 271)

	Abortion Situation														
Psychosocial Characteristic	Mother's life	Rape	Mother's health	Mother's mental health	Mother under 15	Mother unmarried	Family completed	Financial burden	Severe MR	Bedridden for life	Death before 5	CF	Moderate MR	Mild MR	Severe disorder at 40
Religion (non-Catholic)	b	b			b	b	b	a				a			
Infrequent attendance at religious services	b	b	a	a	a	a	a	a	b	a	b	b	a	a	b
Higher education (years)	b	a	b	b		b	b			a	b		b	b	b
Higher income	a	b	a	a	b	b	a	a		b	a	a	a	a	a
Occupation (professional)									a	a	a	a	a	a	
Respondent's gender															
Female				b	a										
Male	a		a			a	a								
Previous elective abortion	b						a	a	b	b	a	a	b	b	
Approval of abortion for CF by spouse	a	a	a	a	a	b	b	a	b	a	a	b	a	b	b
Respondent's siblings	a	a	a	a	a	a	a	a	a	a	a	b	b	a	a
Respondent's mother	b	b	b	b	b	a	a	a	b	b	b	a	b	b	b
Respondent's father	a	a	a	a	b	a	a	a	a	a	a	a	a	a	a

Note: a = associated at zero-order level, $P < .05$; b = associated in stepwise logistic regression, $P < .10$.

TABLE 2—Stepwise Logistic Regression for Abortion for Cystic Fibrosis

Variable	Beta	P	Odds Ratio[a]
Siblings approve abortion for CF (3-point scale)	1.0319	.05	7.88
Spouse approves abortion for CF (3-point scale)	0.9396	.002	6.55
Infrequent attendance at religious services (4-point scale: once a year to weekly)	0.5791	.002	5.68
Spouse would not support carrying a CF fetus to term (3-point scale)	0.7713	.004	4.68
CF doctor approves abortion for CF (3-point scale)	0.3105	.06	1.86

[a] Odds that a respondent at one end of the variable scale will abort, when compared with a respondent at the other end.
Note: n = 222.

Discussion

The major finding, that a minority of parents of children with CF (20%) would abort a CF fetus, suggests that prenatal diagnosis will not lead to a substantial reduction in the number of CF births among this group, at least in the United States. Most parents in this sample apparently do not regard selective abortion for CF as an acceptable reproductive option, although they believe this option should be available to others. Many affected families may not welcome public health efforts to provide prenatal testing for CF.

The majority of CF families may proceed as they did before prenatal diagnosis became available: they will either curtail reproduction (as many do) or take their chances. However, in the wider spectrum of abortion attitudes, the majority would abort for severe MR and many for a child bedridden for life. Optimism about CF was pervasive; families expected their child to reach age 40, hold a full-time job, and marry. With discovery of the gene for CF, optimism has probably increased.

At the other end of the spectrum, few would abort for adult-onset disorders starting at age 40 or 60, for genetic susceptibility to alcoholism, or for sex selection, areas now receiving much attention. Although overall responses suggest that most would not use abortion frivolously, responses to the obesity question aroused our concern. Apparently, there is a not insubstantial minority who would abort for cosmetic purposes if prenatal assessment were possible.

The trimester of pregnancy had less relationship to personal willingness to abort than might be expected. For most fetal conditions, including CF, only 5% fewer women would abort in the second trimester than in the first. This result suggests that newer methods of prenatal diagnosis, such as CVS, which make first-trimester abortions possible, may have little effect on deeply held attitudes.

In genetic counseling, it would be appropriate to acknowledge the effects of extended family networks on patients' views and to encourage patients to discuss the relationship between personal and family values. The attitude of the physician caring for the affected child may influence parents' decisions. Few believed that their CF doctors would support abortion of a fetus with CF. Counselors should discuss these parental perceptions, which may reflect the clinically necessary optimism conveyed by pediatricians, rather than the doctors' actual views. Finally, the relationships between sociodemographic var-

TABLE 3—Stepwise Logistic Regression for Abortion for Severe Mental Retardation (n = 213)

Variable	Beta	P	Odds Ratio[a]
Siblings approve abortion for CF (3-point scale)	0.646	.0003	3.64
Respondent's mother approves abortion for CF (3-point scale)	0.5901	.002	3.26
Had previous therapeutic abortion	1.267	.04	3.55
Infrequent attendance at religious services (4-point scale: once a year to weekly)	0.2029	.097	1.83

[a]Odds that a respondent at one end of the variable scale will abort, when compared with a respondent at the other end.

iables and willingness to abort for fetal conditions reinforce concerns that differential use of selective abortion could make genetic disability a mark of social class and lead to further discrimination. ☐

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Infant Diet and Salmonellosis

Robert L. Haddock, DVM, MPH, Simon N Cousens, MA, DipMathStat, and Catherine C. Guzman, MPH, RD

ABSTRACT

Background: **The Island of Guam has experienced a high incidence of infant salmonellosis for a number of years (age-specific incidence rate of approximately 3700 cases per 100 000 infants in 1984). Interviews of case parents suggested that the use of high-iron infant formula was associated with this problem.**

Methods: **A case-control study was conducted to test this hypothesis. Information on feeding practices and a variety of medical and socioeconomic factors was collected for 78 laboratory-confirmed *Salmonella*-case infants and 167 control infants.**

Results: **Case infants were less likely to have been breast-fed (odds ratio [OR] = 9.15; 95% confidence interval [CI] = 2.71–30.9) and more likely to have been fed infant formula with an iron content of 10 mg/L or greater (OR=2.96; 95% CI = 1.24–7.08) than were control infants.**

Conclusions: **Although the precise means by which infants are most commonly exposed to *Salmonella* bacteria remain obscure, breast-feeding apparently protects against the development of physician-diagnosed clinical salmonellosis, while the feeding of high-iron infant formula has the opposite effect. (*Am J Public Health.* 1991;81:997–1000)**

Introduction

The incidence of nontyphoid salmonellosis on the island of Guam increased from 78.6 cases per 100 000 population in 1975 to a peak of 218 cases per 100 000 in 1984. The incidence has declined to 110 per 100 000 in 1989, still more than 5 times the US rate for the same year.[1] Infants typically account for about 50% of *Salmonella* cases on Guam. Another unusual characteristic of the epidemic is the presence of *S. waycross,* a serotype that is rare elsewhere in the world but accounts for about 30% of all human *Salmonella* isolates on Guam. In 1984, only 3 of 35 862 cases of salmonellosis reported in the United States were due to *S. waycross;*[2] all three cases were residents of Guam who were referred to the state of Hawaii for medical treatment (A. P. Liang, personal communication, 1987).

Several investigations have failed to link infant cases of this epidemic to the consumption of foods with an animal origin (including milk), to secondary spread from infected family members, or to the ownership of pets or domestic animals.[3-5] Infant salmonellosis cases occur with approximately equal frequency in both civilian and military communities as well as in families of both higher and lower economic status.[3] Investigations of infant day care and formula preparation practices have not provided any clues to the origin of infant salmonellosis cases.[4]

During the course of an earlier unpublished study of this problem, we were impressed by an apparent association between infant salmonellosis cases and a history of being fed iron-supplemented brands of infant formula. Samples of these products, obtained both from case households and unopened from grocery shelves, were invariably sterile, however.

In 1984 an article in the veterinary literature reported that weanling pigs administered therapeutic amounts of iron orally developed more severe *Escherichia coli* diarrhea than pigs to which the same amounts of iron were administered by intramuscular injection; the authors theorized that increased amounts of available iron in the intestinal tract promoted the multiplication of otherwise harmless numbers of potentially pathogenic bacteria, thus rendering them better able to cause overt disease.[9]

Since iron-enriched formula is commonly fed to infants, we initiated a study to determine if the iron content of formula could be contributing to the problem of infant salmonellosis on Guam.

Methods

A case-control study was carried out using as cases, infants born on Guam who developed laboratory-confirmed salmonellosis before their first birthday. Cases were identified prospectively as laboratory reports were received from the Guam public health reference laboratory. (All medical laboratories on Guam, including

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Robert L. Haddock and Catherine C. Guzman are with the Department of Public Health and Social Services, Agana, Guam. Simon N. Cousens is with the Department of Epidemiology and Population Sciences, London School of Hygiene and Tropical Medicine, University of London.

Requests for reprints should be sent to Dr. R. L. Haddock, Territorial Epidemiologist, Department of Public Health and Social Services, PO Box 2816, Agana, Guam 96910.

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TABLE 1—Crude odds ratios (OR) for Iron Content of Formula Fed at Onset of Illness in Non-Breast-Fed Cases of Infant Salmonellosis and Non-Breast-Fed Age-Matched Controls

Formula	Cases	Controls	OR*	95% CI
Low-iron	13 (17.3%)	37 (30.0%)	1.00	
High-iron	62 (82.7%)	88 (70.0%)	2.96	1.24–7.08

*OR and 95% CI estimated retaining age-matching.

military facilities, routinely submit isolates of *Salmonella, Shigella,* and *Vibrio* to this laboratory for confirmation and serotyping.)

Controls were identified by selecting the next two births listed in the Guam birth registry following the case-birth registration, thereby matching for age and season of birth. If an infant selected as a control was deceased, or a salmonellosis case, or had previously been selected as a control, the next registered birth was selected. If a selected infant had been treated by a physician for diarrhea or if efforts to locate parents were not successful, the next birth certificate was selected, and the process was repeated until two controls had been enrolled in the study.

Parents of case infants were asked what milk their infant was being fed when the diarrhea developed, and parents of control infants were asked what milk their infant was being fed at the same age. Formula brands were assigned to one of two categories: low iron (less than 10 mg/L) or high iron (10 mg/L or more). Upon completion of each interview, data on the age, ethnicity, education, and marital status of the parents, infant birthweight and APGAR score, complications of pregnancy or delivery, month of pregnancy during which prenatal care began, number of prenatal physician visits, and number of previous births to the mother were excerpted from birth certificates.

Names of all study infants, without identification as to whether they were cases or controls, were submitted to the Special Supplemental Food Program for Women, Infants and Children (WIC) staff to determine if they had [illegible] enrolled in that program.

Statistical Methods

Cross-tabulations and simple statistical tests, including χ^2 tests, were performed using the computer package PC-SAS.[10] Odds ratios (OR) and their 95% confidence limits (CI) were estimated by conditional logistic regression analysis, retaining the matching of controls to cases as suggested by Breslow and Day,[11] using the computer package EGRET.[12]

Results

Of 96 cases of salmonellosis reported in children less than 1 year of age during this study, 18 (18.8%) could not be traced. In selecting controls, birth certificates of 389 infants were examined. Of these, 131 (34%) were excluded from the study for various reasons, 112 being untraceable. Of the remaining 258 controls, 167 had no history of treatment by a physician for diarrhea and were enrolled in the study.

Cases were less likely to be receiving breast milk at the time they developed salmonellosis than their age-matched controls (3.9% vs 25.2%), suggesting that failure to breast-feed infants on Guam is associated with an approximately ninefold increase in the risk of symptomatic salmonellosis infection (OR = 9.15, CI = 2.71–30.9). This association remained after taking into account potentially confounding variables.

To compare infants with regard to the iron content of infant formula received, those infants receiving breast milk were excluded, leaving a total of 75 cases and 125 controls. Cases were more likely than controls to be receiving infant formula with a high iron content (OR = 2.96, CI = 1.24–7.08; Table 1). Ninety-six percent of cases enrolled in the WIC program received a high-iron formula compared with 75% of cases not enrolled. Among controls, 80% of infants in the WIC program received a high-iron formula compared with 67% of those not in the program. No association between education of parents or marital status of mother and type of infant formula used was observed.

Each of the variables included in Table 2 was then considered individually as a potential confounder of the association between type of infant formula and risk for salmonellosis using the technique of conditional logistic regression. Controlling WIC enrollment led to the greatest reduction in the estimate of the odds ratio (OR = 2.54, CI = 1.02–6.28), while controlling marital status led to the greatest increase (OR = 3.85, CI = 1.51–9.84). A conditional logistic regression model including the variables listed in Table 2 was then fitted. Problems of collinearity prevented the simultaneous inclusion in the model of terms for father's ethnic group and age. All other variables were included. When father's ethnic group was omitted from the model, the estimated OR of the association between high-iron infant formulas and salmonellosis was 4.74 (CI = 1.28–17.57). Excluding father's age from the model produced an OR estimate of 4.04 (CI = 1.09–14.94).

Discussion

It has been suspected for some time that infants born on Guam may be infected with *Salmonella* bacteria by some means other than the more commonly recognized modes of transmission. The discovery that both indoor and outdoor environments on Guam were heavily contaminated with *Salmonella* bacteria[13,14] and that the incidence of infant salmonellosis in particular was linked to seasonal changes in environmental conditions[15] seemed to support the hypothesis that infant salmonellosis might frequently be a consequence of contamination of the infant's environment rather than the result of consuming contaminated food products or contact with infected family members.[16] It was also suspected that the immunologic naivete of newborn infants was a key factor in facilitating infection by the relatively small number of *Salmonella* bacteria that were likely to be encountered in an apparently clean home environment and that were harmless to other family members. The results of this study suggest that absence of breast-feeding and high levels of iron in some brands of infant formula may also be factors that contribute to an increased risk for infants' developing clinical salmonellosis on Guam.

Reported complications associated with pregnancy or delivery was the only other variable to show a substantially elevated OR in our univariate analysis (Table 2). This trend remained in both multivariate models (OR = 2.79, CI = 0.76–10.26, father's ethnic group excluded, and OR = 3.68, CI = 1.08–12.55, father's age excluded). Complications reported in more than one case included wrapping of umbilical cord around neck of infant (14), meconium staining of infant (8), Caesarean delivery (6), and premature rupture of membranes (2). These infants may have been at increased risk of being identified as salmonellosis cases because of decreased resistance to infection or because they received more intensive medical attention following delivery. While we did not measure either disease-specific or general infant immunity levels, our analysis did use the number of antenatal physician visits as a proxy measure of the tendency of parents to utilize medical services.

Other possible biases in this study include loss of 18 (18.9%) cases and 11 (28.8%) controls. Factors that contributed to the high rate of unlocateable cases and controls included the mobility of younger families, the concurrent use of as many as three different house-numbering schemes in some areas, and the frequent absence of addresses or telephone numbers in patient records. The fact that case address information obtained from health care providers was more current than control address information listed on birth certificates may have affected the success rates for locating individuals from these two groups.

Of initially selected controls, 91 were excluded from the analysis because they had a history of physician treatment for diarrhea and might therefore have been unidentified cases of salmonellosis. Trial inclusion of these children in the analysis led to some reduction in the final estimate of the OR of the association between type of infant formula and risk of salmonellosis, as might be expected if misclassification of disease status had occurred (OR = 2.34, CI = 1.22–9.16), but would not materially alter our interpretation of the results.

A biased estimate of the OR might also have arisen through failure to control other confounding variables. While socioeconomic status has not been controlled directly in the analysis, maternal and paternal education, both indicators of socioeconomic status, have been controlled. Enrollment in the WIC program, which is restricted to low-income families, has also been controlled. Cost differential between low- and high-iron formulas is not a factor, as both formulations cost the same.

Weaning practices could also be a confounding factor, but since 60% of children in the study were age 3 months or less, it seems unlikely that this could have played a substantial role in the observed associations. Formula preparation practice was not examined because an earlier investigation suggested that this was not a risk factor for infant salmonellosis on Guam.[4]

TABLE 2—Crude odds ratio (OR) for Various Physical, Demographic, and Socioeconomic Variables: Infant Salmonellosis Cases vs Age-Matched Controls

	Cases		Controls			
Variable	No.	Percentage	No.	Percentage	OR	95% CI
Sex						
Males	48	62	97	58	1.00	
Females	30	38	70	42	0.89	0.51–1.57
Birth order						
First	30	38	60	36	1.00	
Later	48	62	106	64	0.90	0.53–1.56
Birth weight (g)						
2500 or more	72	92	157	94	1.00	
<2500	6	8	10	6	1.25	0.43–3.60
Mother's status						
Married	61	78	114	68	1.00	
Unmarried	17	22	53	32	0.62	0.33–1.15
Ethnic group of mother						
Pacific Islander	44	56	90	54	1.00	
Asian	18	23	48	29	0.86	0.44–1.67
Other	16	21	29	17	1.22	0.52–2.84
Ethnic group of father						
Pacific Islander	38	52	56	43	1.00	
Asian	17	25	37	29	0.79	0.38–1.65
Other	16	23	36	28	0.67	0.27–1.62
Complications of pregnancy or delivery						
No	44	56	116	69	1.00	
Yes	34	44	51	31	2.10	1.07–4.15
Hospital of birth						
Civilian	59	76	122	73	1.00	
Military	19	24	45	27	0.87	0.30–2.50
Mother's age (years)						
<20	8	10	23	12	1.00	
20–29	49	63	106	64	1.24	0.52–2.95
≥30	21	27	36	24	1.68	0.65–4.37
Father's age (years)						
<20	4	6	5	4	1.00	
20–29	33	49	69	53	0.63	0.16–2.41
≥30	31	46	55	43	0.73	0.18–2.87
Mother's education						
Secondary	47	60	112	67	1.00	
College	31	40	54	33	1.43	0.81–2.52
Father's education						
Secondary	38	60	76	60	1.00	
College	25	40	50	40	1.07	0.57–2.03
Number of prenatal visits						
1–8	30	38	63	38	1.00	
9–11	24	31	56	34	0.95	0.48–1.89
≥12	24	31	46	28	1.12	0.51–2.44
APGAR score at 1 min						
1–8	53	68	113	68	1.00	
9–10	25	32	52	32	1.03	0.56–1.91
APGAR score at 5 min						
1–8	11	14	21	13	1.00	
9–10	67	86	144	87	1.00	0.41–2.45
WIC client						
No	51	65	130	78	1.00	
Yes	27	35	37	22	1.74	0.94–3.21

Note: OR and 95% CI estimated retaining age-matching.

It has previously been reported that breast-feeding and available dietary or serum iron can influence the course of infections.[17–22] Since mother's milk contains only about 1.5 mg of iron per liter and normal full-term infants exclusively breast-fed for 4 to 6 months by mothers in normal health are unlikely to develop iron deficiency,[23] it may be that maintaining a relatively low concentration of iron in the neonatal gastrointestinal tract is one of nature's strategies for protecting infants from bacterial infections during the critical first few months of life. Commercial iron-supplemented infant formula, on the other hand, contains from 12 to 19 mg of iron per liter and would provide the intestinal flora of infants fed these products with from 8 to 12 times more iron than is available to breast-fed infants.

Although one of the objectives of the Guam WIC program is to encourage breast-feeding, this study found that breast-feeding was less common among WIC program participants than among nonparticipants (19% vs 28% among controls). The enticement of offering free infant formula to new mothers may be more effective in determining actual nursing practice than present educational efforts designed to foster knowledge and attitudes favorable to breast-feeding. Current regulations of the WIC program stipulate that infant formula containing less than 10 mg of iron per liter may be provided only when a physician determines that the recipient infant has an existing medical condition that contraindicates the use of iron-supplemented formula. This study suggests that such a requirement may contribute to the problem of infant salmonellosis on Guam.

For the 10-year period 1981 to 1990, the average annual age-specific incidence rate for infant salmonellosis on Guam peaked at age 2 months (approximately 4580 cases per 100 000) but fell by more than 50% by age 6 months (approximately 2140 cases per 100 000). This result suggests that it would be advisable to breast-feed Guam newborns for at least 6 months and to delay feeding iron-supplemented infant formula for the same period when breast-feeding is not possible.

While we have only studied the relationship between salmonellosis and infant diet, it is possible that other bacterial diseases are similarly affected. It has already been demonstrated using an animal model that oral iron supplementation can cause an exacerbation of diarrhea due to *E. coli* infection,[9] and both pathogenic strains of *E. coli* and *Campylobacter* (a bacteria similar epidemiologically to *Salmonella*) are known to be important causes of childhood diarrhea in many countries.[24–26]

Although a relatively high nonresponse rate among both cases and controls and the other potential sources of bias outlined in this report suggest that a cautious interpretation of our conclusions is warranted, we believe that the data presented here, together with a biologically plausible mechanism previously reported, warrant further investigation. Certainly any inexpensive and readily implemented strategies that have the potential to significantly affect the important worldwide problem of infant diarrheal disease merit serious consideration. □

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Controlled Trial of *Giardia lamblia:* Control Strategies in Day Care Centers

Alfred V. Bartlett, MD, Steven J. Englender, MD, MPH, Betty A. Jarvis, RN, Lois Ludwig, RN, John F. Carlson, and James P. Topping

ABSTRACT

Background: Strategies for control of *Giardia lamblia* in day care differ in numbers of children treated and in costs to parents and day care operators. The effectiveness of these strategies has not been systematically evaluated.

Methods: We conducted a prospective randomized controlled trial comparing three strategies for control of *Giardia* in infant-toddler day care centers: Group 1, exclusion and treatment of symptomatic and asymptomatic infected children; Group 2, exclusion and treatment of symptomatic infection only; Group 3, exclusion and treatment of symptomatic infection, treatment of asymptomatic infection in the center. The study included 31 day care centers with 4180 child-months of observation. *Giardia* prevalence was determined before intervention and 1, 2, 4, and 6 months later; new infants and toddlers were tested on admission.

Results: Initial *Giardia* prevalences were 18% to 22% in the three groups. *Giardia* was identified in 10.5% of 676 new infants and toddlers entering study day care centers during the 6-month follow-up. *Giardia* prevelances by intervention group were 8%, 12%, and 7% at 1 month, and 7%, 8%, and 8% at 6 months.

Conclusions: The stricter intervention resulted in greater cost in terms of child day care and parents' work days lost, but did not result in significantly better control of *Giardia* infections in this day care environment. (*Am J Public Health.* 1991;81:1001–1006)

Introduction

Giardia lamblia infection and diarrhea have been increasingly identified in association with child day care centers, especially those caring for infants and toddlers: day care center outbreaks have been reported from Atlanta, Tucson, Houston, and Phoenix, with 17–54 percent of children in these outbreaks found to be infected.[1–4] A prospective study in Phoenix, Arizona day care centers found *Giardia* to be the most frequently identified enteric pathogen in diarrhea outbreaks and in sporadic diarrhea among infants and toddlers.[4]

Asymptomatic children in day care centers where *Giardia* diarrhea is identified are also frequently infected.[1,2,5] These asymptomatic infections are of theoretical concern in developing disease control strategies because asymptomatic carriers often excrete the infective cyst form of the parasite.[5,6] However, the actual relation of such asymptomatic excretion to *Giardia* transmission in day care centers has not been documented.

For almost 10 years, the Maricopa County (Greater Phoenix) Department of Health Services had utilized a very strict intervention strategy to control giardiasis in centers where the parasite was identified by passive laboratory-based surveillance and case investigation:

- If the case is a day care center child or worker, all children and staff were tested for *Giardia*.
- All symptomatic and asymptomatic *Giardia*-infected children and adults were excluded from the center, with recommendation for treatment.
- Excluded children or staff were readmitted to the center when treatment was completed and two *Giardia*-negative stool examinations had been documented.
- The center was given advice regarding hygienic practices to reduce transmission of enteric microorganisms.

Although this strategy had substantial impact on families and centers involved, its effectiveness in *Giardia* control had not been evaluated. We therefore conducted a controlled study of the effectiveness of this strategy in comparison with other strategies for control of *Giardia lamblia* infection among infants and toddlers in Phoenix day care centers.

Methods

General

This was a prospective randomized (unblinded) controlled trial. All centers having a case of *Giardia lamblia* infection identified in an infant (age 0–11 months) or toddler (age 12–35 months) during the period October 1986–September 1987 were considered for inclusion in the study. Permission of the director and parents was required for participation. On acceptance of participation, each center was assigned to one of the three control strategies, using prenumbered sealed envelopes containing the group assignment derived from a table of random numbers according to a block randomization scheme. If a center de-

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Address reprint requests to Alfred V. Bartlett, MD, Department of International Health, School of Hygiene and Public Health, Johns Hopkins University, 615 North Wolfe Street, Baltimore, MD 21205. Co-authors Englender and Carlson are with the Arizona Department of Health Services; Jarvis, Ludwig, and Topping are with Maricopa County Department of Health Services, Arizona. This paper, submitted to the Journal May 2, 1990, was revised and accepted for publication January 23, 1991.

clined participation, *Giardia* control was managed according to the existing policy and excluded from study and analysis. This study was approved by the Committee on Human Subjects of the Maricopa County Department of Health Services.

Control Strategies (Intervention Groups)

Group 1. Exclusion and recommendation of treatment for symptomatic and asymptomatic infections; readmission after completion of treatment, and two *Giardia*-negative stool examinations by the health department (existing strategy).

Group 2. Exclusion and recommendation of treatment for symptomatic infections only; readmission when asymptomatic, with continued treatment and follow-up testing in the center. No exclusion or treatment of asymptomatic infections.

Group 3. Exclusion and recommendation of treatment for symptomatic infections; readmission when asymptomatic, with continued treatment and follow-up testing in the center. Treatment and follow-up testing of asymptomatic infections in the center, without exclusion.

All groups received the same standard orientation regarding hygiene practices to reduce spread of *Giardia*.

Outcome Variables

Effectiveness of the intervention strategies was assessed by comparing prevalences of *Giardia* among infants and toddlers at 1, 2, 4, and 6 months after intervention, and occurrence of *Giardia*-positive diarrhea among infants and toddlers during intervals between follow-up prevalence test rounds.

Data Collection

Information on each participating center included total enrollment, number of infants (if present) and toddlers, type of organization, operating hours, child group mixing, staff supervising more than one child group, and location within Phoenix (zip code sector). Standardized observations regarding sanitary and child care practices were made by study nurses during the six-month follow-up period in each center. Information on each participating child was obtained at the time of center entry into the study for children enrolled at that time (original enrollees), or on entry to the center for children enrolling during the six-month follow-up period (new enrollees). These data included age, sex, ethnic group, attendance status (regular, drop-in), duration of enrollment (original enrollees) or previous day care (new enrollees), other concurrent day care, parent occupation, household size, presence of siblings by age group, home location (zip code sector), pets, exposure to untreated water (by type), and travel history. Information subject to short-term change was re-ascertained by interview at the time of follow-up test rounds.

Giardia lamblia *Prevalence Testing*

Consistent with existing procedures, *Giardia* infection was determined by examination of a single stool specimen collected by the center or parents and tested in the health department laboratory. Prevalence testing at each specified interval was conducted during a period of two weeks. A minimum testing compliance standard (at least 75 percent of enrolled infants and toddlers) was maintained for all test rounds and, when necessary, the test period was extended until this testing level was obtained. In almost all follow-up prevalence testing rounds this level was exceeded, and the mean proportion of enrolled children tested in follow-up test rounds was 83 percent.

New entry children were initially tested within two weeks of admission. In subsequent test rounds, children enrolling after intervention were retested along with original enrollees. Results of initial *Giardia* tests of original enrollees were known to study nurses, center staff, and parents. Study and center personnel and parents were blinded to results of new enrollee testing and of follow-up test rounds for all children.

Surveillance for Diarrhea

Nurse field workers made daily telephone calls to each study center to identify cases of diarrhea (stools looser and more frequent than normal for that child) among infants and toddlers, and to obtain attendance figures. Visits to each center were made at least weekly to deliver specimen collection materials, reinforce study methods, and verify enrollment. When diarrhea occurred in a study child between follow-up test rounds, a stool specimen was collected for *Giardia* testing. Test results of children with diarrhea were known to study personnel; if *Giardia*-positive, children with diarrhea were managed according to the control strategy for their center.

Parasitologic Methods

Fresh stool specimens were immediately placed in two vials, one containing polyvinyl alcohol (PVA) and the other 10 percent formalin, by the parent or center worker. All specimens were examined by ethyl-acetate-formalin concentration of the formalin-preserved stool, followed by staining with D'Antoni's modified iodine and direct microscopic examination at 100× and 400× magnification. Specimens from children with diarrhea were also examined for *Giardia* trophozoites using trichrome staining of the PVA-preserved stools. For quality control, a subset of specimens was randomly selected and also tested using an indirect immunofluorescent assay for the detection of *Giardia* cysts in feces (Merifluor-*Giardia*, Meridian Diagnostics, Cincinnati, Ohio).

Potential Confounding Variables

Variables which might affect the rate of *Giardia* infection among the children and centers in the different intervention groups were identified from review of pertinent literature. We performed two analyses to assess the association of these variables with *Giardia* infection among Phoenix infants and toddlers, and to examine additional local variables (such as geographic sector of the city or exposure to untreated water used for irrigation of some yards and public parks) which might be associated with infection. The first analysis was a case-control study comparing the 51 new enrollee children without previous day care attendance who were found to be *Giardia*-infected on admission to study centers (cases) with the 482 new enrollee children without previous day care who were not infected on admission (controls). A second analysis compared characteristics of 52 children age 0–35 months who had never attended day care and who were identified as *Giardia*-infected during the study period by health department surveillance, with 52 age-matched controls randomly selected from county birth records who also had never attended day care and who were confirmed as not infected by stool testing. The prevalences of confounding variables identified from the literature or by one of these analyses were then compared in the three intervention groups, to identify those which would need to be considered in analysis of the effect of intervention strategy on *Giardia* prevalence.

Association of center characteristics with prevalence of *Giardia* was examined by comparing the 14 study centers having initial *Giardia* prevalences above the mean for all study centers, with the 17 study centers having lower initial *Giardia* prevalences.

Statistical Analysis

Prevalences of *Giardia* among infants and toddlers in the centers in each of the three intervention groups at each of the specified test points were compared using ANOVA. Comparison of rates and proportions was performed using chi-square, Fisher's exact, and Wilcoxon-Mann-Whitney tests. To examine the effect of intervention group on *Giardia* prevalence, accounting for the effects of confounding variables, a nested ANOVA model was used.

In all of these analyses, the null hypothesis (that no significant association existed between an independent variable and the outcome under study) was rejected if the probability of chance association was <0.05.

Estimated Costs of Control Strategies

Our study was not designed to provide a detailed cost-effectiveness analysis of the various *Giardia* control strategies. However, we did attempt to estimate the costs of these strategies to the families actually participating in the study. Projections regarding the number of children requiring medical treatment, costs of these treatments, and number of day care and parent work days lost were developed for asymptomatic and symptomatic children under the different intervention strategies. These projections were then applied to the actual number of such children in each group at the time of intervention.

The projections assumed exclusion and treatment of all *Giardia*-positive children in accordance with the protocol for their treatment group. In Group 1, duration of exclusion of all *Giardia*-positive children was projected as two weeks or 10 day care days (10 days treatment, four days follow-up testing after treatment). In Groups 2 and 3, duration of exclusion of symptomatic *Giardia*-positive children was projected as one week (five day care days) until end of symptoms and readmission. To develop projections of treatment costs and parent work days lost we conducted a survey which included 38 percent of parents of study children who were *Giardia*-positive at the time of initial intervention.

Results

Centers Entering Study

TABLE 1—Characteristics of Study Day Care Centers, by Intervention Group

Characteristics	Group 1 N = 10 (%)	Group 2 N = 11 (%)	Group 3 N = 10 (%)
For-profit organization	9 (90)	9 (82)	10 (100)
Operate 50–65 hours/week	8 (80)	9 (92)	8 (80)
Accept public agency payment	10 (100)	10 (91)	8 (80)
Accept infants [age < 1 year]	6 (60)	9 (82)	4 (40)
Mean number of infant-toddlers at initial testing	26 (9)*	31 (10)*	30 (11)*
Drop-ins (% of infant-toddlers)	(1)	(<1)	(1)
Mean turnover index [# new infant-toddler enrollees during study + # original infant-toddler enrollees]	0.84 (0.10)*	0.82 (0.14)*	0.70 (0.10)*
Proportion of *Giardia*-positive children symptomatic at intervention	17%	39%	22%

*standard error.

TABLE 2—Prevalences of *Giardia Lamblia* on Initial and Post-intervention Testing among Infants and Toddlers in Study Day Care Centers, by Intervention Group and Enrollment Category

Intervention Group	Enrollment Category	Initial	1 month	2 months	4 months	6 months
1	Already enrolled*	20% (258)	7% (165)	7% (181)	8% (183)	6% (175)
	New**	—	10% (58)	11% (54)	9% (74)	9% (34)
	Total	20% (258)	8% (223)	8% (235)	8% (257)	7% (209)
2	Already enrolled*	18% (308)	12% (193)	14% (211)	12% (217)	6% (207)
	New**	—	12% (32)	6% (67)	8% (93)	15% (60)
	Total	18% (308)	12% (225)	12% (278)	10% (310)	8% (267)
3	Already enrolled*	22% (299)	6% (195)	10% (190)	8% (191)	7% (209)
	New**	—	12% (34)	13% (60)	10% (78)	14% (36)
	Total	22% (299)	7% (229)	11% (250)	8% (269)	8% (245)

*Already enrolled = Children enrolled at initial testing, or enrolled throughout the interval ending in each test round.

**New = Children enrolling during the interval ending in the test round.

During the one-year enrollment period, 35 centers with an index case of *Giardia lamblia* infection in the infant-toddler age group were identified. Of the 32 centers having additional cases of infant-toddler *Giardia* infection, all accepted participation; one center was dropped after intervention for failure to comply with the study protocol and its data were excluded from analysis. The remaining 31 centers completed six months of follow-up for a total 4,180 infant-toddler child-months of observation. Characteristics of these 31 centers are shown in Table 1, by intervention group.

Immediate Observations after Intervention

The overall prevalences of *Giardia* at the time of intervention did not differ significantly among the infants and toddlers in each of the intervention groups (20 percent in Group 1, 18 percent in Group 2, and 22 percent in Group 3). The ranges of *Giardia* prevalences in the individual centers at intervention were 6–32 percent in Group 1, 8–30 percent in Group 2, and 5–47 percent in Group 3.

Table 2 presents overall *Giardia* prevalences by intervention group among children already enrolled (prior to the intervention for initial testing and during the entire interval between tests for each subsequent test round) and for children newly enrolled during each of the intervals between test rounds. Several differences between intervention groups were noted in the first month after intervention. In Groups 1 and 3 there were significant reductions in overall *Giardia* prevalence at one month, to 8 percent and 7 percent, respectively. Although the overall prevalence in Group 2 also declined to 12 percent, the difference was less, consistent with the fact that this intervention did not treat asymptomatic *Giardia*-positive children.

In addition, all groups experienced drops in enrollment after initial identification of *Giardia* and intervention imple-

TABLE 3—Contribution of Original and New Enrollees to the Total Infant Toddler Study Population and to *Giardia Lamblia* Infection at Conclusion of Six Months' Follow-up, by Initial *Giardia* Test Result

Enrollment and Test Result Category	Proportion Six Months Post-Intervention Total Infants and Toddlers (n = 721) %	 Total *Giardia*-positive Infants and Toddlers (n = 56) %
Original Enrollees		
Initially *Giardia*-positive	12	18
Initially *Giardia*-negative	43	18
New Enrollees		
Initially *Giardia*-positive	4	39
Initially *Giardia*-negative	41	25

TABLE 4—Episodes of Diarrhea and *Giardia*-positive Diarrhea Occurring in Study Infants and Toddlers between Follow-up Test Rounds, by Intervention Group

Intervention Group	Child-months between Test Rounds	Total Number of Diarrhea Episodes	Number of *Giardia*-Positive Episodes (%)
1	832	60	6 (10)
2	1,048	183	29 (16)
3	907	81	16 (20)

mentation. There were significant differences among intervention groups in the proportion of initially *Giardia*-positive children who left study day care centers in the first month after intervention: only 7 percent of *Giardia*-positive children dropped out of Group 3, compared to 25 percent in Group 1 and 20 percent in Group 2. The drop-out rates of *Giardia*-negative children in the first month after intervention in the three groups were 15, 14, and 16 percents, respectively.

Finally, compliance with post-treatment testing was significantly reduced in the groups where such testing was not required prior to re-admission: only 71 percent of children treated in Groups 2 and 3 submitted one or more post-treatment clearance specimens. Compliance with treatment may also have been different: of treated children submitting post-treatment specimens, 23 percent in Groups 2 and 3 remained *Giardia*-positive, compared to 4 percent in Group 1.

Six-Month Follow-Up

Prevalence of Giardia lamblia. As shown in Table 2, Groups 1 and 3 had similar overall prevalences of *Giardia* at each of the follow-up test points. Although *Giardia* prevalences in Group 2 were somewhat higher than those in the other two groups at the 1, 2, and 4 month follow-ups, by the six-month follow-up the overall *Giardia* prevalences in the three groups were virtually identical (7, 8, and 8 percents, respectively).

Contribution of new versus original enrollees and initial test status to outcome. Table 3 shows the contribution of initially *Giardia*-infected and uninfected original and new enrollee children to the total study population and to *Giardia* infection at the end of follow-up. Six months after intervention, original enrollees accounted for 55 percent of infants and toddlers and for 36 percent of *Giardia* infection. Children who had been *Giardia*-infected at the time of intervention were over three times more likely to be infected six months later than children originally not infected. The importance of originally *Giardia*-positive children in total infection after six months was actually greatest in Group 1, despite these children having been subject to the strictest intervention and post-treatment testing.

Initial testing of all new infant-toddler enrollees found 10.5 percent to be *Giardia*-positive upon entering study centers; this proportion did not differ significantly among intervention groups (10 percent in Groups 1 and 2, 12 percent in Group 3). At six months after intervention, 57–75 percent of *Giardia*-positive infants and toddlers in each group were children who had enrolled after the intervention. New enrollees who were *Giardia*-positive on admission comprised only 4 percent of the total infant-toddler population at six months post-intervention, but accounted for 39 percent of all *Giardia*-infected children.

Giardia-associated diarrhea between test rounds. Table 4 shows the number of diarrhea episodes identified between test periods in each intervention group, and the proportion of these episodes which were *Giardia*-positive. The overall rates of diarrhea and of *Giardia*-positive diarrhea were significantly higher in Group 2 than in Groups 1 and 3. However, the proportions of diarrhea episodes which were *Giardia*-positive were not significantly different among the three groups. Increased testing in response to a greater number of episodes of diarrhea might have resulted in detection of more "*Giardia*-positive" diarrhea in Group 2. However, the testing and treatment of symptomatic *Giardia*-positive children represents an additional control measure which would have had more effect in Group 2 than in Group 1 or 3.

Potential confounding variables. Five characteristics other than day care attendance were found to be significantly associated with infant-toddler *Giardia* infection in new enrollees without previous day care who entered study day care centers, or in the case-control study of *Giardia*-infected 0–35 month old children not attending day care. Comparison of the prevalences of these characteristics among study children in the three intervention groups found very low prevalences of two (Native American ethnic group, 2–5 percent; well water exposure, 1–2 percent). Two other characteristics were more common, but were almost equally prevalent in all groups (household size ≥ 5 persons, 25–29 percent; presence of a 3–5 year old sibling, 28–32 percent). Only one factor, the study child belonging to the toddler (12–35 month) age group (versus the infant age group), was significantly associated with *Giardia* infection and differed significantly in its distribution between the three intervention groups (78 percent in Group 1, 65 percent in Group 2, and 82 percent in Group 3, $p < 0.01$.). The lower proportion of toddlers in Group 2 would be expected to lower the observed prevalences of *Giardia*, while the higher proportion of toddlers in Group 3 would be expected to raise the observed prevalences of *Giardia* in that group.

Among all new enrollees, previous day care center (but not day care home) attendance was significantly associated with *Giardia* infection, but only 8 percent of new enrollees had previously attended a day care center.

No day care center characteristic was significantly associated with higher prevalence of *Giardia* on initial testing, although this analysis was limited by the small total number of centers.

Multivariate analysis. To examine the effects of the *Giardia* control strategies taking into account the differences in age distribution among the day care centers in the three groups, we employed a series of nested ANOVA models which included *Giardia* prevalence (at one of the specified test points for each model), intervention group, proportion of infants and toddlers tested at that point who were in the toddler age group, and individual day care center (nested within intervention group). In these models, the differences in proportion of toddlers were significantly associated with differences in *Giardia* prevalence at the time of initial testing and at 1, 2, and 4 months after intervention. Intervention strategy was not significantly associated with differences in *Giardia* prevalence at any test point.

Estimated Costs of Control Strategies

Our survey of parents of *Giardia*-infected children at the time of intervention found the mean cost of physician visits associated with this diagnosis to be $34 (range $1–$120) and the mean cost of medication to be $12 (range $2–$60). Forty-two percent of these parents reported missing work in association with their child's exclusion from day care, with a projected mean of 6.6 days for parents in Group 1, and 3.5 days for parents in Groups 2 and 3.

Table 5 presents the projected estimates of costs which result from applying these results to each of the intervention groups at the time of intervention and throughout the six months of study including the intervention. These estimates are based on the actual number of symptomatic and asymptomatic children detected at the time of intervention, and on the additional cases of *Giardia*-positive diarrhea detected during follow-up. For the initial intervention, Group 2 had substantially lower projected costs than Group 1 in terms of both medical costs and of day care and parent work days lost. Group 3 had lost time costs similar to Group 2, but had substantially higher medical costs resulting from the treatment of asymptomatic as well as symptomatic children.

TABLE 5—Estimated Costs of *Giardia* Control Strategies in Study Day Care Centers at the Time of Initial Intervention and Total during Six Months of Observation, by Intervention Group

	Cost at Initial Intervention/Total Cost Over Six Months		
	Group 1	Group 2	Group 3
No. of children excluded	52/58	22/51	18/34
Medical treatments required	52/58	22/51	67/83
DCC days lost	20/580	110/255	90/170
Parent work days lost	114/161	32/75	26/50
Physician costs	$1,768/$1,972	$748/$1,734	$2,278/$2,822
Prescription costs	$624/$696	$264/$612	$804/$996

By the end of six months, the large number of cases of *Giardia*-positive diarrhea in Group 2 resulted in a total number of excluded children, and resulting projected medical costs, similar to those of Group 1. Because of the difference in exclusion requirements between Group 1 and the other two groups, the projected numbers of day care and work days lost remained substantially lower in Groups 2 and 3 throughout the period of study.

Discussion

Strategies for control of specific infectious diseases in child day care must be principally based on concern for the health of the infected or ill child, the probability of transmission of infection to other children or adults, the effect of infection on these other persons, and the preventive or curative control measures available. At the same time, these strategies must be developed and implemented taking into account their impact on the child, the family, employers, and day care providers. Physicians and local health authorities have usually had to develop disease control interventions in the absence of adequate information regarding their effectiveness and impact.[6–8]

Giardia lamblia has been recognized as a problem in day care of diaper age children, and *Giardia* infection has been associated with day care center attendance.[1,2,9,10] In Phoenix and in some other day care environments, asymptomatic *Giardia* infection among infants and toddlers in centers has also been documented to be prevalent.[4,5] These asymptomatic infected children present a special disease control dilemma, since limited data suggest no adverse nutritional impact of such infection among well-nourished immunocompetent children,[5] and no increase in diarrhea rate among such children compared with *Giardia*-negative children in the same settings.[4,5] However, these asymptomatic infected children do represent a potential source of infection to others.

The existence of a very strict health department strategy for control of *Giardia lamblia* in Maricopa County (greater Phoenix) day care centers gave us the opportunity to perform this randomized controlled evaluation of that strategy in comparison with two less strict strategies. Our study was designed as an effectiveness evaluation of these intervention strategies under existing conditions of day care centers, turnover of enrollment, and medical treatment provided by community care providers, rather than an efficacy evaluation in which these variables would be controlled by the investigators. For the same reason, we continued to identify day care centers having *Giardia*-infected infants and toddlers through the existing health department surveillance mechanism and continued the health department policy of identifying *Giardia* through a single stool examination.

Several immediate differences emerged between intervention groups. At one month after intervention, significant reductions in *Giardia* prevalences existed in the two groups of centers where both symptomatic and asymptomatic children were treated; a lesser non-significant reduction in prevalence occurred in the group of centers where only symptomatic children were treated. Part of this immediate reduction in *Giardia* prevalence was attributable to *Giardia*-positive children leaving the study centers. Although we did not investigate whether these *Giardia*-positive children moved to other day care centers and infected other children, molecular epidemiological study of *Shigella*-infected day care center children has identified such transfer of infection to other child groups in response to control measures.[11]

At the end of the six-month follow-up period, no control strategy was associated with significantly lower prevalences of *Giardia* among infants and toddlers in study centers, although the six-month prevalences in all three groups were significantly lower than the prevalences at the time of intervention. Of equal importance is the recognition that the 8 percent overall prevalence of *Giardia* at six months after intervention is essentially the same as the 10.5 percent prevalence of *Giardia* among the infants and toddlers who entered study centers during the follow-up period; the majority of these new enrollees had not previously attended a day care center. *Giardia*-positive new enrollees accounted for almost 40 percent of *Giardia* infection at six months. The unexpectedly high prevalence of *Giardia* among Phoenix infants and toddlers without previous day care histories, and the high turnover of children characteristic of this center population, suggests that *Giardia* control strategies for Phoenix day care centers cannot expect to reduce prevalence below this apparent background level.

These characteristics of the Phoenix population of infant-toddler center users probably limit the external validity of our findings to similar populations. It may be that in an infant-toddler population with very low prevalences of *Giardia* and low rates of center turnover, strict control strategies might result in elimination of *Giardia* from a given center. In the Phoenix day care center environment, however, our data indicate that the existing strategy did not achieve substantially greater control of *Giardia* infection, especially in comparison with the strategy that treated both symptomatic and asymptomatic infection with less stringent exclusion requirements.

Our data do not permit general recommendations regarding the need for *Giardia* testing of young children in day care who experience diarrhea. They do suggest that, in environments comparable to Maricopa County, *Giardia* testing should be considered if microbiological testing is indicated in an episode of diarrhea in an infant or toddler (whether in a day care center or not) or if a day care center is experiencing a diarrhea outbreak.

Our projected estimates of the cost of the different strategies were consistent with parents' and day care owners' experience that the existing strategy resulted in substantial numbers of child day care and parent work days lost. While initially the least costly, Group 2 experienced more diarrhea during follow-up, resulting in total medical costs comparable to the existing strategy.

Our data do not permit distinction of diarrhea caused by *Giardia* infection from diarrhea of other etiology occurring in a child who was coincidentally infected by *Giardia*. However, if the reduced initial costs of Group 2 were accomplished at the expense of additional preventable child illness, this would clearly be an unacceptable strategy. The greatest difference in projected costs between strategies, and that most important to parents and day care operators, resulted from the difference in exclusion requirements. The *Giardia* prevalence results and diarrhea experience in Group 3 suggest that these less stringent requirements did not result in an outcome substantially different from that of the stricter exclusion policy.

Although the cost of conducting controlled trials of public health disease control strategies restricts the usefulness of this approach, documentation to the extent possible of the effects and impact of these strategies can represent an important contribution to day care disease control efforts. Certainly, in various child day care environments additional information is needed on the control of *Giardia lamblia* and other infectious diseases important in preschool age children. ☐

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Characteristics of Maternal Employment during Pregnancy: Effects on Low Birthweight

ABSTRACT

Background: Although maternal employment is considered a risk factor for low birthweight (LBW), the manner in which employment might affect birthweight is poorly understood. In this analysis, selected charateristics of employment during pregnancy were examined for effects on pregnancy outcomes.

Methods: Work characteristics included the number of hours per week, physical activities, and environmental conditions. The outcomes of interest were fetal growth retardation (<2500 grams at term) and preterm delivery (<37 weeks). The study population consisted of 2711 non-Black, married mothers who participated in the 1980 National Natality Survey (NNS). The NNS data were merged with data from the 1977 revision of *The Dictionary of Occupational Titles* (DOT) from which measures of occupational physical activities and environmental exposures were obtained. Logistic regression was used in the analysis.

Results: Those who worked 40 or more hours per week were more likely than women who worked fewer hours to have a low birthweight delivery at ≥37 weeks. No physical or environmental characterics of work were associated with low birthweight or preterm delivery.

Conclusions. Non-Black married American women may face a risk of delivering low-birthweight babies at or near term only if they work 40 or more hours each week. However, the lack of risk associated with other characteristics of work may be a function of measurement error in the DOT data source or of low levels of exposure in the analysis population. (*Am J Public Health*. 1991;81;1007–1012)

Mary D. Peoples-Sheps, RN, DrPH, Earl Siegel, MD, MPH, Chirayath M. Suchindran, PhD, Hideki Origasa, PhD, Andrea Ware, MPH, MEd, and Ali Barakat, MS, PhD

Introduction

Infants weighing 2500 g or less at birth may be small for gestational age, the result of a preterm birth, or both.[1] The two low-birthweight (LBW) categories share many risk factors[2] but are caused by different physiological mechanisms.[3] For women who work during pregnancy, characteristics of employment have been associated inconsistently with both outcomes.

Studies in developing countries suggest that hard physical labor during pregnancy has adverse effects on fetal growth.[4–6] In industrialized nations, standing,[7–10] carrying or lifting heavy loads,[8,11] strenuous physical effort,[12–15] work on assembly lines[12] and industrial machines,[13] and some occupations that involve physical effort, such as chambermaid, janitor, and hospital worker,[8,16] have been associated with delayed fetal growth, preterm birth, or both. However, in other investigations, each of these findings has either not been supported[8,12,17–22] or has not been examined.

Studies of environmental conditions such as temperature, dust, and moisture have found few associations with either category of LBW.[8,17,23] However, one study found loud noises associated with lower birthweight,[8] while in another, an "environment" index was associated with preterm birth.[13]

Physical and environmental exposures at work may influence pregnancy by decreasing uterine blood flow,[24,25] by stimulating production of catecholamines[26,27] or both. Presumably, these processes can be accelerated or decelerated by amount of exposure. Hence, studies in which number of hours worked per day or week are taken into account are potentially informative. However, the studies that have included these variables have produced inconsistent results.[7,8,12,13,18,19,22,28]

Most of the research in this area has been done in Europe,[10–13,19,20,29] with some recent reports from Montreal.[8,16] Of the 10 American studies reported in the literature,[7,9,14,15,17–19,22,28,30] four found no significant effects of maternal employment.[17–19,28] Also, the analysis of the Collaborative Perinatal Project data of the early 1960s[7] may not be applicable to current working populations. Most of the US studies have focused on a few physical activities (notably that of standing) and on hours and duration of work to the exclusion of environmental exposures. Also, confidence in the findings of the studies of American women is tempered by measurement problems. In some cases, employment characteristics were assigned post hoc by the investigators or other experts to occupational titles.[7,9,17,18] Variations by individual or work site were not taken into account, and the validity of the coding schemes was unknown. Another approach was to collect employment data after the delivery,[19,22] when recall may

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Mary D. Peoples-Sheps, Earl Siegel, Chirayath M. Suchindran, and Andrea Ware are with the School of Public Health at the University of North Carolina, Chapel Hill. Hideki Origasa is with R&D Laboratories, Eisai Company, Ltd, Tokyo, Japan, and Ali Barakat is with the An Najah National University, Nablus, West Bank, Israel.

Requests for reprints should be sent to Mary D. Peoples-Sheps, RN, DrPH, Associate Professor of Public Health Nursing and Maternal and Child Health, University of North Carolina, Chapel Hill, NC 27599-7400.

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have been clouded by elapsed time, the relative significance of the birth event, the health of the infant, or all of these factors. Homer et al.[14,30] avoided these sources of error by imputing employment characteristics from an independent data base. However, they may have introduced a new source of self-report bias because the data were collected by surveying workers.

The purpose of the present study was to examine the effects of selected physical and environmental characteristics of maternal employment and amount of work (hours/week) on fetal growth and preterm delivery in a representative sample of contemporary American women. To avoid the measurement error in assessment of job characteristics associated with self-report and expert opinion, data on employment characteristics were obtained from an independent source derived primarily from work-site observations.

Methods

Data Sources

This study is an analysis of the 1980 National Natality Survey (NNS), a follow-back survey of a probability sample of registered births in the United States. Approximately 1 in 400 live births ≥2500 g and 1 in 95 live births <2500 g are in the NNS sample, which totals 9941. Data were collected from birth certificates and requested by mail from hospitals and other health care providers. In addition, detailed sociodemographic and behavioral data as well as basic information on employment during the 12 months before the birth were requested by mail from mothers who were married at the time of delivery. The data collection process involved three mailings and an attempt to interview nonrespondents by telephone. Overall response was 79.5% from married mothers and 76.1% from hospitals. All data were collected retrospectively, within 1 year of the 1980 deliveries.[31,32]

The NNS data set was merged with the 1977 Dictionary of Occupational Titles (DOT), which includes descriptive data on more than 12 000 distinct occupations in more than 75 000 sites.[33] DOT data were collected by on-site interviews with workers and from observations of jobs as they were performed.[34] A detailed job analysis schedule[33] was completed for each job and, if feasible, verified by the worker's immediate supervisor.[34] One or two workers were observed for each job analysis. Scores on 44 work characteristics were then assigned by the analyst based on detailed guidelines and examples.[35]

Study Population

The study population consisted of 2711 non-Black (97% White) singleton live births to married mothers who worked during pregnancy and provided sufficient data on occupation and industry to match with DOT job characteristics. NNS subjects with imputed data entries on employment items, birthweight, or gestational age at delivery were not included in the 2711. The population was limited to married women because employment data were available only from the mothers' questionnaire. This restriction caused the number of Black mothers eligible for analysis to be so small (n = 399) that they were excluded from the analysis.

Two sets of analyses were conducted to distinguish between the effects of work characteristics on fetal growth and on gestational age, since the two conditions are due to different physiologic processes.[3] For analyses in which term low birthweight was the dependent variable, infants who were clearly preterm (<37 weeks) were removed from the analysis population. Similarly, in analyses of preterm births, infants who were low birthweight (fetal growth retarded) at term (≥37 weeks) were removed. The purpose of this was to exclude infants with identifiable morbidities from the groups to which low birthweight and preterm births were compared. In addition, women who started work after the eighth gestational month were excluded from analyses of preterm births.

Derivation of the analysis samples is show schematically in the Appendix. Each exclusion produced more homogeneous populations with lower risk of adverse pregnancy outcomes. In general, the analysis samples were of lower parity, had more education, received more prenatal care, and gained more weight that the exclusion groups.

Variables

The dependent variables, low birthweight (<2500, ≥2500 g) and preterm delivery (<37, ≥37 weeks), were constructed from the birth certificate source in the NNS. Number of hours worked per week in a company, business or organization was taken from the NNS mothers' questionnaire and categorized as 1 to 20, 21 to 39, and ≥40. Three measures of physical activity were constructed from DOT data. Physical strength was derived from field observations of standing, walking, sitting, lifting, carrying, pushing, and pulling objects; jobs were categorized as sedentary, light, medium, heavy, or very heavy based on these activities.[35] The other physical activities were climbing/balancing and stooping/kneeling/crouching/crawling. A score of 1 for presence or 0 for absence was recorded for each DOT job on these variables.[35]

Six DOT characteristics were used as measures of environmental conditions: (1) work performed outdoors without effective protection from the weather; (2) extreme cold with or without abrupt variations in temperature; (3) extreme heat with or without abrupt variations in temperature; (4) constant or intermittent loud noises; (5) physical hazards (e.g., work on scaffolding, exposure to electrical shock, exposure to toxic chemicals); and (6) elements in the atmosphere (i.e., fumes, odors, dusts, mists, gases, or poor ventilation.) These characteristics were also scored 1 or 0, as above.[35]

DOT job titles do not correspond exactly with the NNS census codes for occupation and industry. To merge the two files, a bridge tape created for this purpose was used. Each DOT value on the bridge tape represented the level of the characteristic for incumbents of a given occupation in the 1971 Current Population Survey. Levels for the physical exertion variable were population means, while values for the remaining DOT characteristics were percentages (of the population in the given occupation with the characteristic present).[36,37]

In this analysis, the scores described above were categorized into low, medium, and high groups, based on distribution in the NNS population. For symmetrical distributions, the lowest and highest deciles were used to designate low and high groups. In cases of asymmetrical distributions with long tails caused by large percentages of jobs in the population that did or did not have the particular DOT characteristic, cut points at the tails were used to designate low or high categories.

Six NNS variables were considered potential confounders.[2] Maternal age, parity, education, and adequacy of prenatal care[38] were categorical variables constructed from birth certificate data. Number of cigarettes smoked per day after confirmation of pregnancy was a continuous variable derived from the mothers' questionnaire. Prepregnant weight adjusted for height was based on data from hospital and mothers' questionnaires. To construct the variable, the National Center for Health Statistics (NCHS) formula, weight/height$^{1.5}$ was applied to the population of all working women.[39] Adjusted weights were categorized as lean (≤14 percentile), normal (15 to 84 percentile), and overweight (≥85 percentile).

Pregnancy complications and weight gain were available in the NNS, but these variables were not controlled in the analysis because effects of work on fetal growth and gestational duration may occur through these conditions. Other potential confounders such as family income, wantedness of pregnancy, and interval since last pregnancy were not used because of missing data. Indicators of previous LBW or preterm birth were not available.

Data Quality

Cases with missing or imputed entries for birthweight, gestational age, and all characteristics of employment were excluded from the analysis. Fewer than 1% of missing data were imputed for the other variables, with one exception, parity at 13.1%. The imputation process and other NNS editing and cleanup processes are described elsewhere.[31,32] Birthweight and gestational age combinations were examined for plausibility according to David's criteria.[40] Forty-five subjects had implausible combinations and were excluded from the analysis. NNS occupation and industry combinations were compared with DOT job characteristics to assess face validity. The investigators concluded that correspondence between the two data sets was acceptable.

Analysis

Analysis was performed by multiple logistic regression using sampling weights. Low-birthweight infants were oversampled in the NNS, and there was differential reporting by age and birth order, resulting in unequal probabilities of selection.[41] Sampling weights were assigned to each birth in each stratum, defined by birthweight, age, and birth order. The analyses reported here involved computing weighted means and proportions and then constructing logistic regression models using the weights. The analysis was conducted using the software PROC RTI LOGIT written for fitting models to complex survey data.[42] This procedure incorporates sampling weights into the estimation of the logistic regression parameters.

Several logistic regression models were constructed with low birthweight as well as preterm delivery as dependent variables. In addition to the work characteristics, the six potential confounders were used as independent variables. Some models included interaction terms for the DOT characteristics by number of hours worked and parity. Generally, a *P* value of .05 was used to designate significance of the F tests of each parameter. Odds ratios and 95% confidence intervals (CI) were also calculated.

Findings

The study population (Table 1) had relatively low percentages of low birthweights, preterm births, and maternal risk characteristics. Each of the analysis populations had similar distributions.

In the population who delivered at term (Table 2), the percent LBW increased directly with number of hours worked, but the only DOT variable distributed in the hypothesized direction was extreme cold. When the effects of confounders were controlled, only number of hours worked was statistically significant. With 1 to 20 hours as the standard, the odds of having a LBW infant were 1.4 (CI = 0.80–2.36) times greater for women who worked 21 to 39 hours, and 1.7 (CI = 1.03–2.68) times greater for women who worked ≥40 hours (Table 3).

Among women who started work at 8 or fewer months of gestation and had infants who were not LBW at term (Table 4), there is no discernible pattern in the distribution of preterm births by number of hours worked or by the DOT characteristics, with one exception: extreme cold. None of the employment characteristics was significant when confounding variables were controlled in logistic regression analysis.

Discussion

The findings suggest that working long hours during pregnancy is associated with low-birthweight deliveries at 37 or more weeks' gestation for married, non-Black women. None of the other characteristics of maternal employment are strongly related to low birthweight at term or to preterm delivery in this population.

The relationship between working long hours and delivery of a full-term, low-birthweight infant is difficult to interpret in light of the other, null, results. The relationship held when each of the other characteristics of work was included in the model, suggesting that long working hours has a strong independent effect on low birthweight. Working 40 or more hours per week may reflect elements of psychological stress or physical fatigue that were not represented in the other variables. Unlike the dramatic findings of Mamelle, et al.[13], analysis of interactions between hours worked and the other work characteristics produced no additional information. However, because interaction analyses were sometimes based on few observations in the combined categories of interest, the possibility cannot be ruled out that working long hours in jobs with identifiable characteristics is the influential factor.

TABLE 1—Distribution of Pregnancy Outcomes and Potential Confounders for 2711 Non-Black Married Women Who Worked during Pregnancy

Characteristic	Percentage of the Population[a]
Birthweight (g)	
<2500	4.2
≥2500	95.8
Gestational age (weeks)	
<37	6.8
≥37	93.2
Maternal age (years)	
<20	6.7
20–24	34.8
25–29	37.7
≥30	20.9
Parity	
0	48.6
1–3	48.3
≥4	3.2
Education (years)	
<12	7.7
12	44.6
>12	47.7
Prenatal care	
Adequate	79.8
Less than adequate	20.2
Prepregnant weight standardized by height	
Lean	16.2
Normal	74.8
Overweight	8.9
Number of cigarettes smoked per day since pregnancy confirmed	
0	75.5
1–9	7.3
10–19	8.5
≥20	8.6

[a] Weighted for differential probabilities of selection. The sums of percentages may not add to 100 because of rounding.

Some measurement issues also must be considered in interpreting this finding. One is the retrospective nature of the NNS data. In the year following the 1980 birth, respondents probably would have provided accurate information about occupation and industry, but recall of the number of hours worked per week may

TABLE 2—Distribution of Characteristics of Work in the Analysis Population and Percent Low Birthweight in Each Category, 2379 Infants of Non-Black Women Who Delivered at Term (≥37 weeks) and Worked during Pregnancy

Characteristic of Work	Percentage in the Population[a]	Percentage Weighing <2500 g[a]
Hours worked/week		
1–20	16.5	1.4
21–39	26.1	1.7
≥40	57.4	2.0
Strength		
Low	11.4	2.8
Medium	78.9	1.7
High	9.7	1.9
Climbing/balancing		
Low	41.2	1.7
Medium	47.1	2.2
High	11.7	1.1
Stooping		
Low	14.9	1.7
Medium	75.5	1.9
High	9.6	1.4
Work performed outdoors		
Low	67.6	1.9
Medium	26.6	1.7
High	5.8	1.6
Extreme cold		
Low	84.7	1.8
Medium	10.7	1.7
High	4.6	2.7
Extreme heat		
Low	66.3	1.7
Medium	24.3	2.4
High	9.4	1.8
Loud noises		
Low	32.4	1.8
Medium	63.0	1.9
High	4.6	2.1
Physical hazards		
Low	37.3	1.8
Medium	53.4	1.9
High	9.3	1.9
Atmospheric elements		
Low	65.1	1.9
Medium	24.7	1.9
High	10.2	1.5
Total	100.0	1.9

[a] Weighted for differential probabilities of selection.

have been blurred by time and subsequent events. Also, the number of hours worked per week could have changed during pregnancy, but the NNS questionnaire did not capture this dynamic aspect. Employment data were limited to the mother's chief job. Some women may have worked additional jobs that could have extended the number of hours worked per week. The effects of these deficiencies on the findings would be underestimation of the coefficients of hours worked.

The null findings from analysis of DOT physical and environmental factors should be considered from three perspectives. First, these results are consistent with other studies of American populations,[17–19,22,28] but they contrast sharply with findings from major studies in France[11–13] and Canada.[8,16] Perhaps the working conditions of US women are more favorable for birthweight and length of gestation than those in other industrialized countries, or perhaps the relatively advantaged and homogeneous character of our sample is not comparable to the diverse populations in the French and Canadian studies. Also, the quality of the DOT data (discussed below) may have been inferior to data collected directly, albeit retrospectively, from working women in the French and Canadian studies.

A second possible interpretation is derived from the fact that the low-risk status of the population may have diminished our ability to detect effects of physical and environmental exposures within the sample. Because the study population was sociodemographically advantaged, they may have tolerated physical and environmental stresses of work better than a less advantaged group. Or they may have undertaken jobs with fewer physical demands and more comfortable working environments than less advantaged women. Another analysis of the 1980 NNS data found that advantaged women tended to work in the professional, technical, and managerial sector,[43] where physical and environmental insults are likely to be few. We also found evidence that the study population had some control over work-related exposures. There were considerably higher percentages of women who stopped working in the first 3 months of pregnancy who had jobs categorized as high on each of the DOT factors. This tendency was also found, although not as consistently, for women who worked short (1 to 20) hours.

A third possibility, misclassification bias in DOT data, warrants close attention. The only positive result in this study was based on NNS data, whereas all analyses of DOT variables produced null results. We selected the DOT as our source of employment data because, as an independent source derived primarily from on-site observations, it allowed us to avoid self-report and expert opinion biases in job measurement. The DOT has been used extensively for research[44] and has been shown to correlate reasonably well with another large national employment data base.[45] Linking of macrolevel data bases with individual data is advocated as an informative tool for initial epidemiologic studies. However, an inherent bias toward the null of coefficients derived from multiple regression analysis has been reported.[45] In addition, a number of documentation and procedural problems with data collection and assignment of scores in the DOT have been identified.[34] These problems could have resulted in misclassification of the variables of interest. We attempted to identify misclassification by examining DOT values in each occupation or industry for face validity before any analyses were conducted. Also, after analyses of physical and environmental variables were completed, we created three separate measures of high vs low risk for physical exertion from NNS occupation data only, and repeated the logistic regression analysis with each. Consistent with our DOT findings, none of these measures was associated with either of the outcomes. (The definitions of these variables

TABLE 3—Results of Multiple Logistic Regression Analysis for Non-Black Women Who Worked during Pregnancy and Delivered at 37 or More Weeks' Gestation

Variable	Beta	Standard Error	F[a]	P	OR[b]	95% Confidence Interval
Maternal age (years)						
<20	0.088	0.069	1.59	.207	0.9	0.80–1.05
≥30	0.253	0.056	20.37	.000	1.3	1.15–1.43
Number of cigarettes smoked per day	0.048	0.007	46.51	.000	1.1	1.03–1.06
Prenatal care less than adequate	0.4756	0.176	7.32	.007	1.6	1.14–2.27
Hours worked						
21–39	0.319	0.275	1.35	.245	1.4	0.80–2.36
≥40	0.508	0.245	4.31	.038	1.7	1.03–2.68

[a] F tests are with (1, α) *df*.
[b] OR = odds ratio.
Note: Test for overall model: F 12.60; *df* 6; *P* < 0.0001.

TABLE 4—Distribution of Characteristics of Work in the Analysis Population and Percentage of Preterm Deliveries in Each Category for 2511 Infants of Non-Black Women Who Started Work at 8 or Fewer Months' Gestation[a]

Characteristic of Work	Percentage in the Population[b]	Percentage of <37-Week Deliveries[b]
Hours worked/week		
1–20	16.4	6.4
21–39	26.2	7.6
≥40	57.4	6.9
Strength		
Low	11.4	7.9
Medium	78.7	6.6
High	9.9	8.9
Climbing/balancing		
Low	41.3	7.1
Medium	46.9	6.9
High	11.7	6.9
Stooping		
Low	14.7	6.1
Medium	75.7	7.2
High	9.6	6.9
Work performed outdoors		
Low	67.4	7.1
Medium	26.8	7.2
High	5.6	5.1
Extreme cold		
Low	84.5	6.8
Medium	10.7	6.8
High	4.8	11.1
Extreme heat		
Low	66.2	7.0
Medium	24.4	7.2
High	9.4	6.4
Loud noises		
Low	32.7	8.3
Medium	62.7	6.3
High	4.5	7.1
Physical hazards		
Low	37.2	6.9
Medium	53.7	7.2
High	9.2	6.0
Atmospheric elements		
Low	64.9	6.7
Medium	25.0	8.1
High	10.1	6.0
Total	100.0	7.0

[a] 1 month = approximately 4 weeks.
[b] Weighted for differential probabilities of selection.

and results of the analyses are available from the authors upon request.)

The promise for statistical control of potential confounders in the NNS was substantial, but because of missing data, several critical variables such as income, pregnancy interval, and alcohol consumption were not used. Other key variables (e.g., previous LBW, previous pregnancy complications, psychosocial stress, and physical activity outside of work) were not available. While it is unlikely that control of these factors would have altered the null results, they should be included in future research.

In light of the preceding observations, it is prudent to conclude that non-Black, married American women may face a risk of delivering low-birthweight babies at or near term only if they work 40 or more hours each week. However, the lack of risk associated with other characteristics of work may be a function of measurement error in the DOT data source or of low levels of exposure in this sociodemographically advantaged population. With regard to the potential effects of physical and environmental characteristics in the workplace, this study has taken the logical next step in a progression of epidemiologic investigations that includes retrospective self-reports of job exposures, post hoc assignment of exposures by expert panels, and merging of macrolevel job characteristic files with individual-level data. It is now clear that these less expensive approaches to studying maternal employment and pregnancy outcome are not sufficient. Studies based on high-quality data collected prospectively from the women themselves are imperative. ☐

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APPENDIX—Derivation of Analysis Samples National Natality Survey, 1980

Total NNS Sample	9941
Married Subjects	7825
Singleton Live Births	7514
Mother Workded During Pregnancy	4532
Non-Black Live Births	4133
Matched DOT	3333
Complete Data on Work and Outcome Variables	2711

LBW Analysis Sample		Preterm Birth Analysis Sample	
Term Live Births	2379	All Live Births except Term LBW	2511

Source of Bias in Prenatal Care Utilization Indices: Implications for Evaluating the Medicaid Expansion

Greg R. Alexander, MPH, ScD, Mark E. Tompkins, PhD, Donna J. Petersen, MHS, ScD, and Judith Weiss

ABSTRACT

Background: Recent expansions in eligibility for coverage of prenatal care services by the Medicaid program reflect national initiatives to improve pregnancy outcomes. This study investigates the potential impact that completeness of reporting of prenatal care and gestational age variables and strategies to impute missing data may have on evaluations of the Medicaid expansion.

Methods: This study, examining 15 years of vital record data from a single state and comparing 1 year of data from four mid-Atlantic states, selected single live births to resident mothers for analyses. The "day 15" and the "preceding case" methods were used to impute missing gestational age data.

Results: Considerable temporal and geographic variation was detected in completeness of reporting of variables used to construct prenatal care indices. After imputing values for cases with missing data, the proportion of cases for which adequacy of prenatal care utilization could not be determined ranged from 3% to 24% among the states investigated. For those cases where gestational age data could be imputed, the distribution of prenatal care utilization was not markedly disparate from those cases with complete reporting of gestational age.

Conclusions: The results indicate that variations in reporting, decisions regarding the treatment of missing data, and the choice of the denominator can alter prenatal care utilization percentages and have implications for evaluations of the impact of the recent Medicaid expansion on prenatal care utilization. (*Am J Public Health.* 1991;81:1013–1016)

Introduction

Recent expansions in eligibility for coverage of prenatal care services by the Medicaid program reflect policy initiatives to meet Year 2000 national health objectives related to low birthweight and infant mortality.[1] Increased utilization of prenatal care has been associated with improved pregnancy outcomes.[2–7] In an effort to reduce financial barriers to prenatal care, Medicaid eligibility criteria were raised in 1990 to 133% of the federal poverty level, with state options to extend coverage to 185% of poverty. One focus for evaluation of this policy is the extent to which changes in eligibility for state Medicaid coverage will increase utilization of prenatal care and ultimately decrease low birthweight percentages and infant mortality rates.[8]

Prenatal care utilization is commonly measured by composite indices, the most notable being the Kessner Index of Adequacy of Prenatal Care.[2] This index considers the month prenatal care began, the number of prenatal care visits received, and the gestational age at delivery to categorize prenatal care utilization as adequate, intermediate, or inadequate. The adequate category includes cases with a first trimester initiation of care and prenatal care visits that are more than or equal to a recommended number for a pregnancy of a given duration. More visits are required as gestation increases. Initiation of care in the second trimester or an earlier initiation with less than the number of visits recommended to achieve the adequate category is termed intermediate care. Cases with no prenatal care, initiation of care in the third trimester, or an earlier initiation of care with less than the number of visits stipulated to achieve the intermediate category are labeled inadequate care. Modifications and alternatives to this index have been proposed[3,9–11] and used to investigate the relationship of prenatal care to a variety of pregnancy outcome measures.[2–6,9] Unfortunately, descriptions of these indices have often failed to specify how missing or implausible data should be handled.

The treatment of missing data in the construction of these indices is an important issue. In the Kessner index, cases with missing prenatal care data are classified as having inadequate prenatal care utilization.[2] Further, while various methodologies have been suggested to impute missing gestational age data,[12–15] a convention for the use of imputed gestational age data in the calculation of prenatal care utilization indices has not been established. The date of last menstrual period (LMP), needed to calculate the gestational age interval, has been estimated to be incomplete on approximately 20% of the nation's birth certificates.[12]

State-specific analyses of trends in prenatal care utilization before and after adoption of Medicaid options and comparisons of prenatal care among states with varying eligibility levels may be sub-

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Greg R. Alexander is with the Maternal and Child Health Program, School of Public Health, University of Minnesota. Mark E. Tompkins is with the Department of Government and International Studies, University of South Carolina. Donna J. Petersen is with the Department of Health, State of Minnesota. Judith Weiss is with the Department of Health Policy and Management, The Johns Hopkins University.

Requests for reprints should be sent to Dr. Greg R. Alexander, Maternal and Child Health, University of Minnesota, PO Box 197 Mayo, 420 Delaware Street, SE, Minneapolis, MN 55455.

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TABLE 1—Percentage of Missing Components of Prenatal Care Utilization Index by 5-Year Groups: 1974–1988 South Carolina Resident Single Live Births

Missing Component	1974–1978	1979–1983	1984–1988
		Percent	
Trimester that prenatal care began[a]	5.99	0.74	1.10
No. of prenatal care visits[b]	1.09	0.62	1.04
Gestational age interval[c]			
No imputation	16.33	5.31	6.91
Day 15	2.09	1.97	2.21
Preceding case	2.04	1.92	2.15
Total missing[d]			
No imputation	19.05	5.37	7.03
Day 15	7.50	2.21	2.63
Preceding case	7.47	2.18	2.58

[a] Includes cases with a not-stated month prenatal care began or with a stated but nonspecific response, e.g., an actual month (January through December).

[b] Includes cases with a not-stated number of visits or with stated visits of 50 or more.

[c] Includes cases with an incomplete date of last menstrual period or a missing gestational age interval value after imputation, where indicated. Gestational age intervals <20 or >50 weeks are treated as missing.

[d] Includes cases with a missing value for any of the prenatal care utilization index component variables, i.e., trimester prenatal care began, number of prenatal care visits, and gestational age interval from last menstrual period.

TABLE 2—Prenatal Care Utilization Percentages by Gestational Age Imputation Method and 5-Year Groups: 1974–1988 South Carolina Resident Single Live Births

Prenatal Care Groups	1974–1978 (N = 237 347)	1979–1983 (N = 251 247)	1984–1988 (N = 256 175)
No imputation			
Adequate	48.36	55.81	53.61
Intermediate	24.43	29.23	28.45
Inadequate/missing	27.21	14.97	17.94
Missing[a]	19.05	5.37	7.03
Inadequate	8.17	9.59	10.91
Day 15			
Adequate	54.27	57.42	55.88
Intermediate	28.46	30.34	29.96
Inadequate/missing	17.27	12.23	14.16
Missing[a]	7.50	2.21	2.63
Inadequate	9.77	10.02	11.53
Preceding case			
Adequate	54.27	57.43	55.89
Intermediate	28.48	30.36	29.98
Inadequate/missing	17.25	12.21	14.13
Missing[a]	7.47	2.18	2.58
Inadequate	−9.78	−10.03	−11.55

[a] Includes cases with a missing value for any of the prenatal care utilization index component variables, i.e., trimester prenatal care began, number of prenatal care visits, and gestational age (after imputation where indicated).

Note: Because of rounding, columns may not total 100%.

ject to bias due to changes in the completeness of reporting of the variables used to construct these indices and in methodologies employed to impute missing data. This study investigates the potential impact that completeness of reporting of prenatal care and gestational age variables and strategies to impute missing data may have on evaluations of the Medicaid expansion. Two study designs have been employed: (1) examining 15 years of data from a single state, and (2) comparing 1 year of data from four eastern states.

Data and Methods

Computerized South Carolina live birth files for the years 1974 to 1988 were used for the first component of this investigation. For the second component, additional data from three mid-Atlantic states with varying levels of completeness of reporting of gestational age were drawn from the Department of Health and Human Services (DHHS) Region III Perinatal Information Consortium's 1987 Live Birth File.[16] Single births to resident mothers were selected for both analyses.

Gestational age was calculated as the interval from the date of last normal menses to the date of birth and was truncated into completed weeks.[17] Two strategies were employed for imputing missing gestational age data.[15] The "day 15" method imputes a value of 15 for the day of last normal menses for those births for which this single piece of information prevents the calculation of the gestational age interval. The "preceding case" method imputes gestational age by inserting the calculated gestational age interval value of a preceding birth with a corresponding month of LMP, 500-g birthweight interval, and race of mother.

Results

Table 1 displays the proportion of South Carolina cases with missing prenatal care and gestational age information for three 5-year periods. Nearly 6% of the cases in the 1974 to 1978 period exhibited missing or nonspecific data on month or trimester of pregnancy that prenatal care began. Missing gestational age data were found for more than 16% of the cases in the 1974 to 1978 period, but dropped markedly in later periods. After imputation methods were employed to recapture cases with incomplete dates of LMP, gestational age was still missing for approximately 2% of the cases in each period.

Table 2 presents percentages of prenatal care utilization employing three different treatments of missing gestational age data (no imputation, day 15, and preceding case). Based on no imputation, adequate prenatal care utilization was lowest in the 1974 to 1978 period (48.36%). When cases with missing data were combined with the inadequate category, the percentages based on no imputation varied across the three periods from 27.21% to 14.97% to 17.94%. The separate inadequate category increased from 8.17% in the 1974 to 1978 period to 10.91% for 1984 to 1988. There is little appreciable difference in the prenatal care distribution percentages produced by the two alternative methods for imputing missing gestational age values.

The next phase of the analysis utilized 1987 data from four states. The preceding case method was selected to impute missing gestational age values. Table 3 depicts by state the percentages of cases with missing gestational age data, before and after imputation, and missing prenatal care initiation and visit data. States B and C revealed a relatively higher proportion of missing gestational age data, 20.10% and 17.58%, respectively. While imputation markedly reduced this proportion in state C (5.73%), it had little impact in state B, where most cases with missing gestational age data were missing the entire date of LMP, not just the day value.

Prenatal care utilization percentages by state and gestational age imputation group are displayed in Table 4. These percentages are calculated in two different ways. In part A, percentages are based on a denominator of all live births with a separate category for cases with missing prenatal care or gestational age values after the application of the indicated gestational age imputation method. In part B, after imputation where indicated, cases with missing prenatal care and gestational age data are excluded, and percentages are based on a denominator of cases with the prerequisite in-range data elements needed to compute the index.

For states A and B, the imputation of gestational age had little appreciable effect on prenatal care percentages based on all cases (Table 4, part A). However, the proportion of total missing cases was noticeably reduced by imputing gestational age in the other two states. When calculations only included cases with in-range or imputed data values (Table 4, part B), the use of the gestational age imputation method had little appreciable impact on the prenatal care utilization percentages. This result suggests that the prenatal care utilization distribution for those cases for which a missing gestational age value could be imputed was only slightly less adequate than that of cases with complete gestational age data. For example, in state C, where gestational age was imputed for more than 10% of the cases, the adequate percentage was 71.55% before imputation and 71.06% afterward.

TABLE 3—Percentage of Missing Components of Prenatal Care Utilization Index by State: 1987 Resident Single Live Births

Missing Component	State A	State B	State C	SC[a]
Trimester that prenatal care began[b]	4.96	4.73	2.84	0.88
No. of prenatal care visits[c]	2.12	6.68	2.94	1.17
Gestational age interval[d]				
No imputation	5.10	20.10	17.58	7.26
Preceding case	4.69	20.01	5.73	2.16

[a]SC = South Carolina

[b]Missing trimester in pregnancy prenatal care began category includes cases with a not-stated month prenatal care began.

[c]Missing number of prenatal care visits category includes cases with a not-stated number of visits or with stated visits of 50 or greater.

[d]Missing gestational age category includes cases with an incomplete date of last menstrual period or a missing gestational age interval value after imputation, where indicated. Gestational age intervals of <20 or >50 weeks are treated as missing.

TABLE 4—Prenatal Care Utilization Percentages by Imputation Method and State: 1987 Resident Single Live Births

	State A	State B	State C	SC[a]
Part A: Percentages Based on All Cases				
No imputation	9929	70 835	158 468	51 580
Adequate	46.40	57.17	58.52	52.49
Intermediate	32.51	14.71	17.23	28.71
Inadequate	14.57	3.99	6.03	11.50
Missing[a]	6.52	24.13	18.21	7.30
Preceding case	9929	70 835	158 468	51 580
Adequate	46.57	57.22	65.90	55.02
Intermediate	32.68	14.73	19.87	30.29
Inadequate	14.60	3.99	6.97	12.14
Missing[a]	6.14	24.06	7.25	2.55
Part B: Percentages Based on Cases with a Calculated Prenatal Care Index Value (Missing Excluded)				
No imputation	9282	53 740	129 817	47 813
Adequate	49.63	75.36	71.55	56.62
Intermediate	34.78	19.39	21.06	30.97
Inadequate	15.59	5.25	7.39	12.41
Preceding case	9319	53 894	146 974	50 263
Adequate	49.62	75.34	71.06	56.46
Intermediate	34.82	19.40	21.42	31.08
Inadequate	15.56	5.26	7.52	12.46

[a]SC = South Carolina

[b]Includes cases with a missing value for trimester prenatal care began, number of prenatal care visits, or gestational age (after imputation where indicated).

Note: Because of rounding, columns may not total 100%.

Discussion

These data indicate that variations in reporting, decisions regarding the treatment of missing data, and choice of denominator can alter prenatal care utilization percentages. The true potential impact of the Medicaid expansion on prenatal care utilization may be obscured due to these factors. Therefore, evaluations of the Medicaid expansion should include a thorough examination of reporting completeness, a disclosure of strategies employed for imputing missing data, and a description of the methods used to calculate prenatal care percentages.

These data further suggest that it is unwise to categorize cases with missing prenatal care or gestational age data in the inadequate care group. The distribution of prenatal care utilization was not markedly disparate for those cases with imputed gestational age data and those cases with complete reporting of gestational age. This result argues both for employing gestational age imputation strategies and not assuming that cases with missing data had inadequate prenatal care utilization. While the prenatal care utilization of those cases with missing data remains unknown, estimates of inadequate prenatal care usage that include cases with missing data may be inflated.

Whether prenatal care percentages should be calculated from all cases or only those for which the index can be generated is illustrated by data from state B. Based on all cases, 57% of women delivering in that state were placed in the adequate category. Excluding cases with missing data from the calculation resulted in a placement of 75% of women into the adequate category. The ease with which prenatal care percentages can be manipulated into adequate and other than adequate categories should not be ignored, given the political context of an impending evaluation of a major public policy initiative. If we assume that cases for which the index can be calculated are roughly similar or only marginally better in their prenatal care utilization patterns than those without the prerequisite data, the 75% figure may be a better estimate of the proportion of all births in the state with adequate prenatal care utilization. To alternatively assume that most cases with missing data have less than adequate utilization of prenatal care is unrealistic in areas with large proportions of missing data. While cases with missing gestational age data have been shown to display higher risks of poor pregnancy outcomes,[15] all cases with missing data will not necessarily exhibit adverse outcomes or less than adequate prenatal care utilization. Care should be taken in the analysis and reporting of these percentages to ensure that large missing group percentages are not used to represent individuals with less than adequate prenatal care utilization.

The choice between the day 15 or the preceding case method to impute missing gestational age data appears to have little impact on prenatal care utilization distributions. While the day 15 method has been noted to inflate preterm percentages in comparison with the preceding case method,[12,15] the amount of bias introduced by this approach appears insufficient to produce a prenatal care utilization distribution markedly different from that obtained by the preceding case method.

Prenatal care utilization indices are a useful but crude surveillance tool for assessing needs and evaluating programs and policy. Evaluation of the Medicaid expansion by these indices is a heuristic attempt to provide accountability for a major policy initiative that may enhance maternal and infant health in ways not easily appraised by available health measurement tools. Because the components of these indices are subject to reporting bias, and because important characteristics of prenatal care (e.g., quality, comprehensiveness, and continuity) are not considered, it is imperative that these indices be prudently employed in evaluation efforts and be continually scrutinized for bias.

Credible data for programmatic efforts and policy discourse require both accurate and complete reporting and logical and systematic analysis. The three component variables of prenatal care utilization indices contain potentially large measurement error. Vagaries in analytical methods used to compute these indices can further add to the uncertainty regarding the proposed relation between prenatal care utilization and pregnancy outcome and the possible impact of policy initiatives. These methodological concerns require ongoing attention, but they should not be used to divert attention from critically needed efforts to ensure that all women have access to and are encouraged to use prenatal care. □

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Uncompensated Hospital Care for Pregnancy and Childbirth Cases

Terrell W. Zollinger, DrPH, Robert M. Saywell, Jr, PhD, MPH, and David K. W. Chu, PhD

ABSTRACT

Background: The large number of medically indigent patients in the United States is a major concern to policymakers and may be due to recent increases in the number of uninsured people. The purpose of this study was to identify the factors that affect the amount of unpaid hospital charges for services provided to pregnant women.

Methods: Individual and hospital data were collected on a representative set of 235 pregnancy and childbirth patients with unpaid hospital charges from 28 hospitals in the state of Indiana.

Results: Most of these patients did not have insurance coverage (63.8%), yet the majority were employed in the public or private sector (72.3%). Over half (55.5%) of the total uncompensated care amount for this group was from the $1000 to 2499 debt category. The median charge for these patients was $1468, of which the typical hospital was able to collect only 25.5%.

Conclusions: The findings support the belief that any national effort to expand the availability of health insurance coverage to women through increased employment will not totally eliminate the uncompensated care problem. The findings also indicate that rural hospitals face the uncompensated care problem mainly because a significant portion of rural patients are without adequate health insurance coverage. (*Am J Public Health.* 1991;81:1017–1022)

Introduction

The large number of medically indigent patients in the United States, a major concern to policymakers, can be attributed to increases in the number of uninsured and underinsured people. This increase in indigent patients, especially pregnant women, has created additional financial burdens for hospitals. In our earlier studies examining uncompensated care patients in hospitals,[1,2] we found that pregnancy and childbirth accounted for the largest proportion of uncompensated care cases (19.1%) and the largest percentage of uncompensated dollar amounts (17.4%). Researchers have found that a lack of adequate prenatal care leads to low birthweight in infants, which in turn contributes to a higher incidence of infant mortality.[3–7] Other research indicates that younger women who have lower incomes and less education, reside in rural areas, or are Medicaid recipients tend to have less access to prenatal care.[3,6,8–11] Until recently, the eligibility requirements for Medicaid had been increasing, and as a result, relatively fewer women and children were being covered by Medicaid.[3,12–15] Denied access to prenatal care increases the hospital cost because of higher frequency of subsequent complications in delivery and perinatal care, which leads to even higher amounts of unpaid hospital bills.[4,7]

In this study, patients belonging to the pregnancy and childbirth subset of the original sample of uncompensated care inpatients[1] were examined to identify the factors that affect the amount of unpaid hospital charges for services provided to pregnant women. The identification of these factors could lead to the development of policies designed to remove existing patient, provider, and system barriers to prenatal care[3,4–6,16–18] and facilitate medical care access for pregnant women.

Methods

Uncompensated care normally consists of bad debt, charity care, contractual allowances, and courtesy discounts. The uncompensated hospital care costs considered in this study were only those originating from the bad debt and charity care categories. Medicare and Medicaid patients were also excluded because bad debts arising from their failure to pay deductibles are reimbursed by the Medicare and Medicaid programs. Bad debts are amounts due from patients who are expected to pay, but for some reason do not. Charity care is care that is provided to patients who, prior to admission, are generally identified as unable to pay and who are not expected to pay. The hospitals in this study did not consistently classify patients as "bad debt" or "charity care" cases.

The sampling scheme used in this study selected a representative set of Indiana hospitals as well as a representative set of patients from those hospitals with unpaid hospital charges. First, a stratified

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Terrell W. Zollinger and Robert M. Saywell are with the School of Public and Environmental Affairs of Indiana University. David K. W. Chu, who was at Indiana University at the time of this study, is now with the College of the Holy Cross, Worcester, Mass.

Requests for reprints should be sent to Terrell W. Zollinger, DrPH, Indiana University, School of Public and Environmental Affairs, 801 W Michigan Street, Indianapolis, IN 46202-5152.

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TABLE 1—Total Number and Percentage of Uncompensated Care Pregnancy/Childbirth Cases by Dollar Category

	N	Total Sample, %	0–$499, %	$500–$999, %	$1000–$2499, %	$2500–$4999, %	$5000–$9999, %	χ^2 *P* Value
Total uncompensated care cases	235	100.0	22.6	15.7	50.2	10.4	1.3	
Location								
Urban	150	63.8	69.8	59.5	62.7	62.5	66.7	.875
Rural	85	36.2	30.2	40.5	37.3	37.5	33.3	
Hospital size								
Large	98	41.7	49.1	43.2	37.3	41.7	66.7	.611
Medium	39	16.6	17.0	10.8	20.3	8.3	0.0	
Small	98	41.7	34.0	45.9	42.4	50.0	33.3	
% Originating from hospital's county	174	79.1	78.0	81.8	80.7	66.7	100.0	.583

random sample based on urban/rural classification of 27 of the 115 acute care general hospitals in Indiana was drawn. Additionally, a tertiary teaching hospital was purposely included in the study to represent that class of hospital. Of the 28 hospitals, 14 (50%) were classified as urban, 16 (57.1%) as small (fewer than 150 beds), 5 (17.9%) as medium (150 to 399 beds), and 7 (25%) as large. All large hospitals were located in urban areas. Second, once the sample of hospitals was identified, a systematic random sample of those patients who did not pay all of their hospital bill during the 1986 calendar year was selected by using a 1/20 sampling fraction. This sampling scheme identified 235 (13.9%) patients hospitalized with the primary diagnosis of pregnancy and childbirth who received uncompensated care. Total revenue for the study hospitals was compiled from the *1986 Indiana State Board of Health Financial Disclosure Report.*[19]

The pregnancy and childbirth cases receiving uncompensated hospital care in our sample were classified by urban/rural hospital location on the basis of whether the hospital was located within a Standard Metropolitan Statistical Area and by amount of the unpaid hospital bill of the patient.

The data were analyzed with the Student's *t* test and analysis of variance to test for significant differences in means; tests for independence were assessed with the chi-square test. The criterion for statistical significance was $P < .05$.

A multiple regression model was constructed by using the percentage of the charges left unpaid as the dependent variable. Potential explanatory variables were included in a stepwise-forward procedure (using $P < 0.10$ for inclusion) to determine which variables had a significant impact on the dependent variable, controlling for the other variables. Since the interval-ratio scale variables used in the model were skewed, log transformed versions of these variables were used in the model to improve their normality.

The results of this study should be interpreted cautiously because of the inherent limitations of the study data. This study was partly based on secondary patient data reported to the hospital by individual patients upon hospital admission. Also, the employment data did not distinguish between part-time and full-time employment, they merely indicated whether a person was reported as being employed or not.

Results

As shown in Table 1, 61.9% of the 235 cases had unpaid hospital charges of $1000 or more. Fewer than one fourth of the cases (22.6%) left unpaid charges under $500. Nearly two thirds (63.8%) were treated in urban hospitals with 41.7% treated in large hospitals, another 41.7% in small hospitals, and only 16.6% treated in medium-sized facilities. On average, 79.1% of the pregnancy and childbirth patients with unpaid bills originated from the county in which the hospital was located. There were no statistically signincant relationships between the distribution of the unpaid amount and the location of the hospital, the size of the hospital, and the number of uncompensated care patients originating from the same county as the hospital.

The study group received a total of $311 382 in uncompensated care (Table 2). Individual uncompensated care amounts ranged from $4 to $7975. Less than 5% of the total uncompensated care amount was from debts smaller than $499; over half (55.5%) of the total uncompensated care amount was from the $1000 to 2499 debt category. Patients in this study with unpaid hospital charges of $5000 or more were responsible for only 6.4% of the uncompensated care.

The total hospital charges for these patients was $469 999, of which study hospitals were only able to collect an average 25.5%. The median total hospital charges for these cases was $1468 with a median uncollected amount of $1226 (83.5%). Patient characteristics, as summarized in Table 3, reveal that the typical pregnancy and childbirth patient with unpaid hospital charges was 23.3 years old, was married (61.8%), spent 3.2 days in the hospital, and was discharged to self/home care (84.7%). The age and marital status of the patients showed no significant relationships with the amount of the hospital bill left unpaid. As the length of stay increased, so did the amount of the hospital bill left unpaid. In addition, a larger proportion of the patients with greater amounts of unpaid hospital bills were discharged to long-term health care facilities. Those with insurance coverage tended to have fewer unpaid hospital charges, and those without insurance tended to leave larger amounts unpaid.

The majority (77.6%) of the study patients reported to be employed in either the public or private sector or self-employed, yet only 36.2% had some form of health insurance coverage. The unemployed group left significantly higher amounts of their hospital bill unpaid than the patients who were employed.

A second analysis evaluated patient and hospital variables by dollar amount of uncompensated care as shown in Table 4. Nearly two thirds (63.2%) of the uncompensated amount was reported by urban hospitals. However, the mean uncompensated care dollar amounts per case did not

TABLE 2—Selected Financial Characteristics of Uncompensated Care Pregnancy/Childbirth Patients

	Total Sample	Less than $500	$500–$999	$1000–$2499	$2500–$4999	$5000–or More
Total uncompensated care (% of total)	$311 382	$14 901 (4.8%)	$28 439 (9.1%)	$172 688 (55.5%)	$75 401 (24.2%)	$19 953 (6.4%)
Total charged (% of total)	$469 999	$87 003 (18.5%)	$52 381 (11.1%)	$193 430 (41.1%)	$117 232 (24.5%)	$19 953 (4.2%)
Average percentage of charges that is collected	25.5%	76.5%	25.3%	6.9%	7.9%	0.0%
Median total charge to uncompensated care patients	$1468	$1286	$961	$1480	$3211	$6535
Median uncollected amount	$1226	$297	$800	$1365	$2970	$6535

TABLE 3—Uncompensated Care Pregnancy/Childbirth Patients by Selected Demographic Characteristics

	N	Total Sample, %	Less than $500, %	$500–$999, %	$1000–$2499, %	$2500–$4999, %	$5000 or More	*P* Value
Mean age (years)		23.3	23.7	22.4	23.4	23.6	22.0	.734[a]
Distribution of age, y								
0–17	22	9.4	9.4	8.1	9.3	12.5	0.0	.819[b]
18–39	213	90.6	90.6	91.9	90.7	87.5	100.0	
Marital status								
Married	126	61.8	67.3	66.7	56.9	66.7	50.0	
Single adult	76	37.3	32.7	33.3	41.2	33.3	50.0	.761[b]
Widowed	2	1.0	0.0	0.0	2.0	0.0	0.0	
Average length of stay (days)		3.2	3.0	2.5	3.0	4.7	9.3	.000[a]
Discharge status								
Self/home care	199	84.7	92.5	83.8	86.4	66.7	33.3	
SNF/ICF	35	14.9	7.5	16.2	12.7	33.3	66.7	.004[b]
Dead	1	0.4	0.0	0.0	0.8	0.0	0.0	
Primary insurance								
Self-pay	150	63.8	17.0	75.7	83.1	58.3	33.3	.000[b]
Commercial and Blue Cross	63	26.4	71.7	16.2	10.2	25.0	0.0	
Other	23	9.8	11.3	8.1	6.8	16.7	66.7	
Employment status								
Unemployed	42	22.3	6.7	15.6	31.2	27.8	33.3	
Public and private sector	136	72.3	93.3	75.0	61.1	72.2	66.7	.0150[b]
Self-employed	10	5.3	0.0	9.4	7.8	0.0	0.0	

Note: Because of rounding, some percentage numbers do not add up to 100.
[a] *t* test for mean values.
[b] χ^2 test.

differ significantly for the rural and urban hospitals. Large hospitals incurred 40.8% of the uncompensated care dollars; medium-sized hospitals, 16.1%; and small hospitals, 43.1%. There was no significant difference in the mean uncollected amounts between the three classes of hospital size.

Self-pay patients accounted for 71.0% of the dollar amount of unpaid hospital bills, compared with 16.4% for commercial and Blue Cross patients. Self-pay patients also accounted for significantly larger mean uncollected amount (along with the "other insurance" category). Of the total uncompensated amount, only 26.5% was generated by those who were unemployed while over two thirds (68.7%) was generated by people employed in the private or public sector. However, there was not a significant difference between the employment groups in the mean amounts of uncompensated care. Married patients accounted for 61.3% of the uncompensated dollar amount, but the mean uncompensated amounts did not differ significantly between the marital status groups.

Table 5 presents uncompensated care information by hospital location. Of the cases treated in urban hospitals, most (65.3%) were in large facilities. Slightly more urban than rural hospital patients resided in the county the hospital was located in, but the differences were not statistically significant.

The study subjects from the urban and rural hospitals did not differ significantly with respect to age, marital status, average length of stay, or employment status. However, cases with unpaid hospital charges reported from rural hospitals were more likely to be self-insured (74.1% vs 58.1%) and more likely to be discharged to self/home care (95.3% vs 78.7%). The median charge for pregnancy and childbirth care for the group who did not pay all of their hospital charges was somewhat higher for those treated in urban hospitals. However, the mean uncollected amount did not differ between the urban and rural hospitals because of the higher collection rates of the urban hospitals.

The multiple regression procedure indicated that two insurance categories: "Commercial and Blue Cross" and "Other" (i.e., HMO, PPO) contributed the

TABLE 4—Amount and Percentage of Total Uncompensated Care Dollars for Pregnancy/Childbirth Patients

	Amount of Uncompensated Care	Percentage of Total	Mean Amount of Uncompensated Care	ANOVA P Value
Location				.815
Urban	$198 924	63.2	$1313	
Rural	$114 458	36.8	$1346	
Hospital size				.866
Large	$127 001	40.8	$1296	
Medium	$50 232	16.1	$1288	
Small	$134 149	43.1	$1369	
Primary insurance				.000
Self-pay	$221 111	71.0	$1474	
Commercial and Blue Cross	$50 920	16.4	$821	
Other	$39 351	12.6	$1711	
Employment status				.266
Unemployed	$84 728	26.5	$1541	
Private and public sector	$168 131	68.7	$1236	
Self-employed	$11 759	4.8	$1176	
Marital status				.972
Married	$162 007	61.3	$1286	
Single	$99 341	37.6	$1307	
Widowed	$2 853	1.1	$1426	

TABLE 5—Selected Patient and Hospital Characteristics by Location and Hospital Size for Pregnancy/Childbirth Patients

	Location		
	Urban	Rural	P Value
Uncompensated care cases, %	63.8	36.2	
Patients by hospital size, %			.000[a]
Large	65.3	0.0	
Medium	26.0	0.0	
Small	8.7	100.0	
Originating from hospital county, %	81.5	75.3	.272[a]
Average age, y	23.0	23.8	.257[b]
Marital status			.671[a]
Married	61.1	63.6	
Single adults	37.6	36.4	
Widowed	1.3	0.0	
Primary insurance			.0153[a]
Self-pay	58.0	74.1	
Commercial and Blue Cross	28.7	22.4	
Other	13.3	3.6	
Average length of stay (days)	3.3	3.0	.405[b]
Discharge status, %			.000[a]
Self/home care	78.7	95.3	
SNF/ICF	21.3	3.5	
Dead	0.0	1.2	
Employed, %	64.7	74.1	.135[a]
Total amount of uncompensated care	$198 924	$114 458	
Total amount charged	$326 231	$143 768	
Average percentage of charged amount that is collected	29.7	18.1	.016[b]
Median total charge to uncompensated care patients	$1555	$1276	.059[b]
Median uncollected amount	$1188	$1230	.815[b]

[a] χ^2.
[b] t test on mean values.

greatest amount to the model. The amount of the total charge to the patient was also significant at the $P < .10$ level. These three variables accounted for nearly 40% of the variation in the percentage of the hospital charges left unpaid. For all three variables, the partial regression coefficients were negative; this indicated that if the patient had some form of insurance, the percentage of the bill left unpaid was lower, and if the total charge was higher, the percentage left unpaid was also lower. Demographic characteristics of the patients (age, marital status, employment) were not found to be important factors in the model. Neither were the discharge status and the length of stay. The two hospital characteristics, urban/rural status and total revenue, were also not important factors in the model.

Discussion

The study of pregnancy and childbirth patients who do not pay all of their hospital bill resulted in three findings that suggest a need to further examine existing public policy as it affects these women: (1) the majority (63.8%) of the pregnancy and childbirth patients who left part or all of their hospital bill unpaid did not have insurance; (2) pregnancy and childbirth patients tended to have overall larger dollar amounts of bad debts relative to other uncompensated care patients; and (3) the collection experience was lower for pregnancy and childbirth patients than for patients from all diagnosis groups as previously reported.[1,2]

The data indicate that the patient's insurance coverage was the major factor in a facility's ability to recover its total charges. Only a third of the patients in this study had any insurance coverage. It is remarkable that 63.8% of the pregnancy and childbirth patients with unpaid hospital charges had no insurance coverage as compared with 44.6% who had no insurance in the sample of patients from all diagnosis groups found in our earlier studies.[1,2] This is a high proportion of uninsured patients in light of the fact that over three fourths of the pregnancy and childbirth patients were reportedly employed. There are three possible explanations: (1) some of the employers do not offer health insurance benefits to their employees, (2) the individual cannot afford or for other reasons chooses not to purchase the insurance, (3) or the insurance policies do not cover pregnancy and childbirth-related hospital services (there may be a waiting period for maternal benefits to take effect). Some of the insurance plans may have high coinsurance or deductibles for pregnancy-related health care, and those patients with insurance coverage may be unable or unwilling to pay their portion of the hospital bill.

These findings, although based solely on the findings of one state, may support the belief that any national effort to expand the availability of health insurance coverage to women through employers will not totally eliminate the uncompensated care problem for maternal and child health patients and the hospitals providing the care. The fact that the majority of the women in this group were employed (77.6%) points to a different set of problems. A policy expanding availability should be considered a step in the right direction to reduce the amount of uncompensated care provided by hospitals. Yet the greater problem may be that employed women in this group cannot afford the insurance premiums and thus remain uninsured or underinsured. On the other hand, if high coinsurance amounts and large deductibles are problems for those with insurance (36.2%), changes in the shared portion would also be required to reduce the amount of the hospital charges determined to be the responsibility of the patients themselves (i.e., lower copayment or deductible amounts).

A second major contrast between the results found in this study of pregnancy and childbirth patients and our previous study of all patients was that pregnancy and childbirth patients tended to have larger total uncompensated dollar amounts. Fewer than 5% of the pregnancy and childbirth cases had uncompensated care amounts of less than $500 compared with 39.2% in the sample of patients from all diagnosis groups. Over one half of the pregnancy and childbirth cases had uncompensated care amounts between $1000 and $2499, compared with only 29.2% in the overall study. This clearly indicates that pregnancy and childbirth patients represent a serious financial problem to the hospital.

The third finding shows that of the total charges to pregnancy and childbirth patients, only a small portion was eventually collected (25.2% compared with 38.8% overall). This points to an apparent inability and/or unwillingness of the facility to seek recovery of these outstanding balances. Hospital administrators may perceive that aggressively pursuing these unpaid bills may be futile and/or not cost-effective on an individual basis, even though collectively such cases represent a large proportion of the hospital's total uncompensated care. The indirect cost of collecting these debts, in the form of poor publicity generated in the community, may be seen as somewhat counterproductive to the mission of the facility.

Although the median dollar amount charged by rural facilities was lower than that charged by urban facilities, the collection rate was also lower, thus the median uncollected amount was approximately the same. These findings indicate that rural hospitals face the uncompensated care problem mainly because a significant portion of rural patients are without adequate health insurance coverage. If this is the case, there is a need to expand insurance coverage to the rural population to ensure that this population has access to comprehensive health care.

The patterns seen with uncompensated hospital care may be related to the overall problem of inadequate prenatal care. Some of the patients who leave hospital bills unpaid may also have failed to obtain adequate prenatal care. The major factor thought common to both problems involves financial barriers to care. Thus hospital administrators are faced with the difficult question of how aggressively they should seek payment for services rendered. Aggressive collection efforts may discourage the medically indigent, higher risk women from seeking prenatal care. However, if adequate prenatal care is not provided by the hospital (even at a loss), high-risk women are more likely to have medical complications. Thus, when these women are finally admitted into the facility, the result could be an even greater amount of uncompensated care than if they had received adequate prenatal care initially. Evidence shows that better prenatal care would be obtained if the financial barriers for patients were removed.[3,6,8–11]

Increasing health insurance coverage for the patients, and not just increasing the availability of health insurance for both those employed and unemployed, is a possible solution. However, an increase in insurance coverage entails more costly public policy interventions. One public policy option could include requiring mandatory insurance coverage to be provided (not merely made available) by employers. However, this option may result in higher cost of goods and services, making American businesses even less competitive in the international arena than they already are. A second option would be to increase Medicaid coverage and eligibility. This option would result in higher taxes and/or the shifting of resources from other public programs into health care programs. These possible consequences must be considered before any policy is enacted. However, we need to recognize that presently a large part of the burden of providing prenatal and hospital care to the indigent population falls on hospitals and only more indirectly on business or the taxpayer. If inroads are to be made into solving the infant mortality problems, policies regarding equitable compensation to hospitals must also be addressed. Further research is needed to examine the demographics and collection experiences of all hospital patients so that the characteristics of the paying and the nonpaying patients can be distinguished. ☐

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Report Cites Vaccine Shortage in Health Centers Serving Poor Children

In 1990, the year that there were to be no more than 500 cases of measles nationally under goals set in 1979 by the Surgeon General, more than 25 000 cases of measles were reported. As community health centers grappled with shortages of government-supplied vaccines, according to a new Children's Defense Fund report, *Vaccine Supply and Low Income Children: Barriers Faced by Community and Migrant Health Center Patients*, more children died from measles in 1990 than in any year since 1971.

"Measles is a particularly sensitive and early indicator of the nation's eroding childhood immunization system because the disease is so contagious," said Sara Rosenbaum, director of CDF's health division. "Other diseases such as mumps and pertussis are showing similar surges. The current epidemic of utterly preventable childhood illness and death must be halted. We have only our lack of collective national will to blame for this crisis."

Yet in the midst of the worst measles outbreak in more than 20 years, the CDF report finds that federally funded community health centers that provide primary health care to nearly 3 million poor children have been experiencing shortages of government-supplied vaccines critical to immunizing poor children against childhood diseases. As a result, the youngest health center patients, all of which live in medically underserved urban and rural communities, run a high risk of inadequate vaccination.

Rosenbaum recommended a number of steps to alleviate the crisis in vaccine shortages and immunization delivery. One step is the immediate expansion of the Centers for Disease Control's (CDC) vaccine contract purchasing program by $80 million.

Rosenbaum said: "Currently CDC simply does not have enough publicly purchased vaccines to meet the needs of both state and local health agencies and health centers. All of these clinics must be assured an adequate supply of vaccines at the lowest available price. The CDC purchase and distribution contract system must be expanded at least to meet the needs of all publicly funded health care providers rather than forcing those providers to stop vaccinations or to purchase on the private market at four times the cost."

Moreover, according to the report, numerous health centers reported that state Medicaid agencies were not reimbursing them for the cost of vaccines administered to Medicaid-enrolled children which they were forced to purchase privately when their health department supplies were exhausted. Others reported that some states refused to cover certain vaccines at all despite federal mandates.

Rosenbaum pointed out that the National Vaccine Advisory Committee in January concluded that low vaccination rates and disease outbreaks are chiefly the result of systemic breakdowns, not parents' willful or negligent failure to seek vaccinations.

The CDF report reviews a number of systemic problems that have led to the growing vaccine crisis, including soaring vaccine costs over the past decade, the increasing number of private physicians who no longer vaccinate children in their offices, increased childhood poverty, and growth in the number of children with no health insurance.

"The childhood immunization crisis is a stark reminder of the severe shortage of basic health services that persist in countless medically underserved urban and rural communities," Rosenbaum said.

If recent trends continue, by the end of the century US immunization rates for infants and toddlers will lag behind those of dozens of Third World countries. Among the 16 countries with higher immunization rates than ours in 1988 to 1989 (as measured by 1-year-olds fully immunized against polio) were Bulgaria, Hungary, Greece, Brazil, China, Mexico, North Korea, Chile, and Rumania. Among the 55 countries with infant polio immunization rates higher than the United States non-White rate were Albania, Cuba, Jordan, Tunisia, Colombia, Iran, Malawi, Botswana, Maritius, Sri Lanka, Mongolia, Jamaica, and Rwanda.

Sociodemographic Factors and the Quality of Prenatal Care

Mary Jo Hansell, BSN, DrPH

ABSTRACT

Background: In this study, maternal sociodemographic factors are examined in relationship to the quality of prenatal health services US women receive.

Methods: Data from the 1980 National Natality Survey and 1980 Fetal Mortality Survey were used for the analysis. Indicator variables for prenatal care quality are the percentages of prenatal visits at which blood pressure and urine were tested, the performance of hemoglobin or hematocrit tests, and the presence or absence of advice regarding salt restriction and diuretics usage during pregnancy.

Results: Distribution of the basic examinations in prenatal care vary according to marital status, parity, education, and residence in a metropolitan or nonmetropolitan county. The advice received concerning salt and diuretics usage was also influenced by sociodemographic variables.

Conclusions: The analyses reveal that prenatal care is not of even minimally acceptable quality for many women. (*Am J Public Health.* 1991;81:1023–1028)

Introduction

Numerous efforts have been made to assess the adequacy of prenatal care obtained by pregnant women in the United States. The adequacy of prenatal care can be measured by examining two components: (1) the timing and quantity of visits and (2) the quality of health services. With little exception to date,[1] studies of the adequacy of prenatal care have concentrated on the first component. Researchers have found that about one third of pregnant women obtain too few prenatal visits or start prenatal care too late. The groups getting the most inadequate care are teenagers, minorities, the least educated, the unmarried, and the poor[2–5] (P. J. Placek, unpublished manuscript, May 15, 1987; a more extensive report on this topic was prepared for the Institute of Medicine by Paul Placek, PhD in his private capacity, with appropriate clearances, and no official support or endorsement by his employer is intended or inferred.) Based upon data from the 1980 National Natality Survey (NNS) and the 1980 National Fetal Mortality Survey (NFMS), this research adds measures of quality to the study of the adequacy of prenatal care in the United States.

Women have varying degrees of control over the timing and quantity of prenatal health services they obtain, because sociodemographic attributes and health beliefs affect their utilization patterns. (When asked about barriers to obtaining adequate prenatal care, 39% of the Medicaid recipients and uninsured women interviewed by the US General Accounting Office cited attitudinal barriers, e.g., not wanting to tell others, being afraid of the pregnancy, and being afraid of the medical tests.)[6] Once a woman has chosen her physician, control over the quality of her health care shifts almost completely to the provider. Theoretically, the woman's sociodemographic attributes should not affect the quality of the services received.

The specific research questions to be answered are whether prenatal care recipients obtain basic examinations recommended by the American College of Obstetrics and Gynecology (ACOG), whether women receive appropriate medical advice concerning the use of salt and diuretics during pregnancy, and whether variability in the quality measures (surveillance and advice) corresponds to maternal sociodemographic characteristics.

ACOG recommends that all pregnant women receive both blood pressure examinations and urine tests, for protein and sugar, at every visit and that hematocrit and hemoglobin levels be assessed at least once during prenatal care.[7] Actual practices are compared to these standards of care.

Although attitudes about salt use in pregnancy had changed over time, by 1977 the ACOG had set the standard that normal women should use salt at the level preferred.[7,8] In other words, physicians should not normally advise salt restriction. Administration of diuretics during pregnancy may be acceptable treatment of underlying hypertensive disease but is not approved for normal pregnancy or for preeclampsia.[9,10] The kinds of advice

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At the time this study was conducted, Mary Jo Hansell was affiliated with the School of Public Health of the University of California, Los Angeles.

Requests for reprints should be sent to Dr. Mary J. Hansell, Assistant Professor, Columbia University School of Nursing, 617 West 168th Street, New York, NY 10032.

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TABLE 1—Explanatory Variables Included in the Regression Models[a]

Variables	Explanation
Maternal sociodemographic factors	
Age	Number of years of age at delivery
Non-Hispanic White[b]	Non-Hispanic White racial-ethnic group
Black[b]	Black racial group
Native American[b]	Native American Indian racial-ethnic group
Asian/Pacific Islander[b]	Asian or Pacific Islander racial-ethnic group
Education	Number of years of education with final value of 17 or more years[c]
Graduate	17 or more years of education
Married	Married at delivery
Parity	Number parity at start of current pregnancy
Metropolitan	Lived in metropolitan county at time of delivery
Prenatal care (PC) quantity and timing measures[d]	
Month PC begun	Month of pregnancy that first prenatal visit made
Number visits	Number of prenatal visits with final value of 30 or more visits[e]
Top visits	30 or more visits made
Maternal health status factors[f]	
High BPs %	Percent of prenatal visits at which blood pressures were above 140/90
Any complications	One or more of a list of complications were noted by the physician
Hypertension	Physician-identified hypertension
Preeclampsia	Physician-identified preeclampsia
CV–renal disease	Physician-identified cardiovascular–renal disease
Prepregnancy weight	Physician-reported maternal weight before pregnancy

[a] Variables are dichotomous (0 = no; 1 = yes), except as indicated.
[b] Compared with Hispanic White (the omitted category).
[c] The NNS and NFMS collapsed the top values to 17 or more years.
[d] Included as controls in the regressions of blood testing, advice about salt restriction, and advice about diuretics usage.
[e] The NNS and NFMS collapsed the top values to 30 or more visits.
[f] Included as controls in the regressions of advice about salt restriction and diuretics usage.

given by physicians regarding salt restriction and diuretics usage are compared with these expert recommendations.

Taffel and Keppel,[11] using the 1980 NNS, examined advice given by physicians about weight gain. They found that the majority of married mothers were given no weight gain limit or were advised to gain less than the amount currently recommended by nutritional experts. The advice varied according to race, educational background, parity, and prepregnancy weight.

Methods

Data

The National Center for Health Statistics based the NNS and NFMS on a complex random sample of 9941 live births and 6386 late fetal deaths that occurred in the United States during 1980.[12] Late fetal deaths are defined as fetal deaths with gestations of 28 weeks or more, or delivery weights of 1000 g or more. To adjust for the oversampling of fetal deaths, deflated sampling weights were attached to all observations for this research, resulting in a nationally representative sample. (Deflated sampling weights were calculated as the original sample weight times the sample size divided by the sum of weights.) The 2.3% of women who made no prenatal visits were excluded from these analyses.

Unimputed data were used from the live birth and fetal death records and from the follow-back questionnaires to the providers of prenatal care and the administrators of the hospitals where deliveries occurred. The prenatal care providers and the hospital administrators (hereafter referred to as the medical source) were asked to refer to medical records to answer detailed questionnaires about the prenatal care given to the women. The NNS and NFMS data have been described in detail elsewhere.[12,13]

Missing medical data may be due to the medical record being incomplete or invalid or it may be that the survey respondent did not fully transcribe the medical record. Donabedian argued that medical record keeping is one dimension of the quality of performance as well as the medium for evaluating other dimensions of care.[14] Also, two sources were asked to transcribe the medical information. Because of these factors, missing data were counted as missing examinations.

Variables

For the entire prenatal period, the medical sources recorded whether hematocrit and hemoglobin levels were obtained. For each prenatal visit, the medical sources recorded blood pressure and result of the urine test for protein. These data were transformed into the percentage of prenatal visits at which blood pressure was measured and the percentage of prenatal visits at which urine was checked for protein. The medical sources were also asked whether physicians advised women to restrict salt or to use diuretics during pregnancy.

The percentage of visits at which urine was tested is skewed to both ends of the range. Therefore, the variable was transformed into a dichotomous variable: whether urine was tested at no visits or at some percentage of visits greater than zero.

Table 1 lists the explanatory variables used in the analyses. The variables, "graduate" and "top visits" are indicator variables representing the top values of the "education" and "number visits" variables. "Graduate" and "top visits" were tested for an effect in each model but were only left in if they improved the model fit (Table 1).

The month of pregnancy that prenatal care was begun and the number of visits that the woman made were entered as controls in the blood test and advice models (Table 1). A provider does not expect to check hemoglobin and hematocrit levels or give particular pieces of advice more than a couple of times and may wait until

TABLE 2—Distribution of Prenatal Care Quality Variables

	Percentage Women
% of visits at which blood pressure was recorded	
None	12.9
1–24	3.5
25–99	16.4
100	67.2
% of visits at which urine tests were recorded	
None	33.7
1–49	6.3
50–99	35.2
100	24.8
Hemoglobin or hematocrit	
Not tested	6.6
Tested	93.4
Salt restriction	
Advised	14.9
Not advised	85.1
Diuretics usage	
Advised	2.6
Not advised	97.4

TABLE 3—Multiple Linear Regression of the Percentage of Visits at Which Blood Pressures Were Recorded on Maternal Sociodemographic Factors

Covariate	Regression Coefficient	Standard Error	Significance Level
Age	0.086	0.070	
Education	0.687	0.146	***
Non-Hispanic White	9.781	2.124	***
Black	−0.168	2.236	
Native American	1.615	3.683	
Asian/Pacific Islander	9.544	3.171	**
Married	4.040	0.918	***
Parity	−1.106	0.235	***
Metropolitan	−1.748	0.623	**

Note: F value = 44.029***; weighted frequency = 13 915. Covariates explained in Table 1.
**$P < .01$.
***$P < .001$.

TABLE 4—Multiple Logistic Regression of Urine Testing for Protein and Sugar on Maternal Sociodemographic Factors

Covariate	Regression Coefficient	Standard Error	Significance Level
Age	0.007	0.004	
Education	0.049	0.010	***
Graduate	0.061	0.108	
Non-Hispanic White	0.644	0.121	***
Black	0.070	0.127	
Native American	0.044	0.209	
Asian/Pacific Islander	0.456	0.187	*
Married	0.225	0.054	***
Parity	−0.071	0.014	***
Metropolitan	−0.065	0.038	*

Note: Likelihood ratio model χ^2 = 425.36 ***; weighted frequency = 14 155; urine never tested (0) = 4480, urine tested (1) = 9675. Covariates explained in Table 1.
*$P < .05$.
***$P < .001$.

after the first visit to do so. If the woman does not return for further visits after the first, or skips the visit when the test or counseling was scheduled, she may be less likely to receive these services at all.

The maternal health status factors listed in Table 1 were included as controls in the advice models because of their possible relationship with the sociodemographic and advice variables.

Statistics

Multiple linear regression analysis as implemented on the Statistical Analysis System (SAS)[15] was done in order to determine the statistical significance of sociodemographic factors in explaining the level of blood pressure monitoring during prenatal care. Logistic regression analysis using SAS[16] was used to estimate the effect of the sociodemographic factors on the occurrence of urine testing, blood testing, and the giving of appropriate advice.

Results

Prenatal Surveillance

During 1980, nearly one third of the women who obtained prenatal care (32.8%) did not receive blood pressure examinations at all of their prenatal visits. Providers skipped all blood pressure examinations on 12.9% of prenatal care recipients (Table 2). Over three quarters of women did not have urine tests for protein at all of the visits they made. One third of prenatal care recipients had no urine tests done during prenatal visits (Table 2). Prenatal care providers obtained no blood test for hemoglobin or hematocrit on 6.6% of the women who were prenatal care recipients (Table 2).

The set of sociodemographic factors (i.e., age, education, race or ethnicity, marital status, parity, and county of residence) explains only a minimal amount of the variance in the percentage of visits at which blood pressures were taken (R^2 = .027).[17] Nonetheless, most of the sociodemographic factors are statistically significant in the regression. For each additional year of education, there is an increase of two thirds of a percent in the number of visits at which blood pressure was taken. Non-Hispanic Whites and Asians and Pacific Islanders received blood pressure exams at almost 10% more of their visits than Hispanic White women (the excluded group). Native American and Black women did not differ in a statistically significant way from Hispanic White women. Married women had blood pressures taken at 4% more of their prenatal visits than unmarried women. For each additional level of parity, women had blood pressure checked at 1% fewer of the visits they made. Finally, women living in metropolitan counties had blood pressures taken at almost 2% fewer of the visits than did women living in nonmetropolitan counties (Table 3).

Turning to the urine test regression, the model χ^2 statistic shows that the set of sociodemographic variables are useful in classifying women in the test or no-test groups ($P < .001$).[17] The maximum likelihood estimate χ^2 statistic for the individual parameters tests the hypothesis that the parameter equals zero. According to this test, women with higher education

TABLE 5—Multiple Logistic Regression of Blood Testing for Hemoglobin (Hgb) or Hematocrit (Hct) on Maternal Sociodemographic Factors

Covariate	Regression Coefficient	Standard Error	Significance Level
Age	0.018	0.008	*
Education	0.013	0.183	
Graduate	1.349	0.393	***
Non-Hispanic White	0.211	0.224	
Black	−0.001	0.233	
Native American	−0.295	0.343	
Asian/Pacific Islander	0.563	0.433	
Married	0.552	0.094	***
Parity	−0.114	0.025	***
Metropolitan	−0.123	0.074	

Note: Likelihood ratio model χ^2 = 266.59***; weighted frequency = 14 155; Hgb or Hct not tested (0) = 869, Hgb or Hct tested (1) = 14 155. Control variables are month prenatal care began, number visits, and top visits. Covariates are explained in Table 1.

*$P < .05$.

***$P < .001$.

TABLE 6—Multiple Logistic Regression of the Advice Regarding Salt Restriction Given by the Physician on Maternal Factors

Covariate	Regression Coefficient	Standard Error	Significance Level
Age	−0.000	0.397	
Education	0.015	0.017	
Graduate	−0.155	0.161	
Non-Hispanic White	0.253	0.280	
Black	0.160	0.294	
Native American	−0.209	0.470	
Asian/Pacific Islander	0.007	0.386	
Married	0.297	0.109	**
Parity	−0.058	0.026	*
Metropolitan	−0.222	0.062	***

Note: Likelihood ratio model χ^2 = 364.99***; weighted frequency = 7577; salt restricted (1) = 1407, not restricted (0) = 6169. Control variables are month prenatal care began, number visits, top visits, hypertension, preeclampsia, % high BPs, and pre-pregnant wt. Covariates are explained in Table 1.

*$P < .05$.

**$P < .01$.

***$P < .001$.

were more likely than women with less education to receive urine tests at their prenatal visits. This was also true for non-Hispanic Whites and Asians or Pacific Islanders relative to Hispanic Whites. Married women were more likely to be tested at visits than unmarried women. Lower parity women were more likely than high-parity women to receive urine testing. And again, nonmetropolitan dwellers were at an advantage over metropolitan dwellers (Table 4).

As shown in Table 5, the sociodemographic variables are useful in classifying women as recipients of the blood test for hemoglobin or hematocrit ($P < .001$). Of the independent factors, maternal age, a graduate education, marital status, and parity are statistically significant. The likelihood that a woman would receive a blood test increased with age, as it did for married women and women of lower parity. Women with 17 or more years of education received blood tests for hemoglobin or hematocrit significantly more often than less educated women (Table 5).

Advice Concerning Salt Restriction and Diuretics Usage

Physicians reported advising salt restriction for 34.5% of hypertensive women, 36.7% of preeclamptic women, and 13.4% of women who were free of any physician-identified complications. Fifteen percent of the overall sample were advised to restrict salt (Table 2).

As shown in Table 6, the set of sociodemographic, health status, and prenatal care factors is useful in classifying women according to the kind of advice that they received regarding salt restriction ($P < .001$). The sociodemographic attributes that decrease the log-likelihood of being advised to restrict salt are being unmarried, having higher parity, and residing in a metropolitan county.

The final logistic regression shows that the set of sociodemographic and control variables is useful in classifying women according to whether they received advice to use diuretics ($P < .001$). Physicians were less likely to advise the use of diuretics for women who had higher parity, were younger, more educated, unmarried, or living in metropolitan counties (Table 7).

Discussion

Although these data are somewhat dated, it was deemed essential to use a nationally representative sample to analyze this health policy topic. Prenatal care data from the National Maternal and Infant Health Survey, the more recent version of the NNS and NFMS, will not be available for analysis until early 1992.

These results point to the importance of including quality of care measures in examining the adequacy of prenatal care. Prenatal care is far from being of even minimally acceptable quality for many women who have made the effort to engage the prenatal care system. High proportions of women were not tested, or not regularly tested, for the basics in prenatal care: blood pressure, urine, hemoglobin, and hematocrit. The advice that many women received concerning salt restriction and the use of diuretics was not congruent with current practice.

Distribution of the tests done to monitor the health of pregnant women varied substantially according to sociodemographic factors, as did the kinds of prenatal advice given. However, the groups of women who were likely to receive poor surveillance were mainly different from the groups likely to receive poor advice. These inconsistencies mean that physicians who provided prenatal care that was of good quality in terms of basic surveillance frequently gave advice that did not follow current recommendations.

The explanation for these inconsistencies may lie in the large percentage of variation in quality that is left unexplained by these models (Table 3). The setting of the health care practice is one variable that could be important. For example, physicians practicing in academic medical centers may have been aware of and followed new advice guidelines before physicians

TABLE 7—Multiple Logistic Regression of the Advice Given by the Physician Regarding Diuretic Use on Maternal Predictors

Covariate	Regression Coefficient	Standard Error	Significance Level
Age	0.034	0.017	*
Education	−0.147	0.039	***
Graduate	−1.320	0.794	
Non-Hispanic White	0.457	0.779	
Black	0.383	0.813	
Native American	−6.385	21.748	
Asian/Pacific Islander	0.361	1.069	
Married	0.916	0.292	**
Parity	−0.220	0.067	**
Metropolitan	−0.336	0.156	*

Note: Likelihood ratio model χ^2 = 318.86***; weighted frequency = 7384; advised to use drug (1) = 203, not advised to use (0) = 7181. Control variables are month prenatal care began, top visits, any complications, hypertension, cardiovascular–renal disease, preeclampsia, % high BPs, and prepregnancy wt. Covariates are explained in Table 1.

*P < .05.

**P < .01.

***P < .001.

in private practice. At the same time, those busy centers may have been settings where surveillance was missed. Without downplaying these inconsistencies, it seems best to concentrate most upon the surveillance results. The theories about optimal behavior in the areas of salt and diuretics usage had undergone recent revision in 1980. Well-qualified physicians may have known, but been unconvinced, that the recent changes in recommendations about salt and diuretics usage were sound. As pointed out by Hemminki,[1] disregarding changes in medical guidelines may in fact serve the best interests of the patient. She gives the example of x-rays, which were recommended during pregnancy and then found to be dangerous and rescinded.

The variability in quality of care according to maternal sociodemographic attributes may be explained in two ways. First, physicians allow their clinical judgment to be significantly affected by the nonmedical characteristics of the woman they are treating. Second, women of different sociodemographic backgrounds are obtaining health services from providers of substantially different quality. Although these data contain no information on the provider of prenatal care, we know that unmarried women would have higher likelihood of being on Medicaid.

Overall, the groups found to be in a relatively deprived position in terms of surveillance are already known to be the groups who obtain care later in pregnancy and in apparently inadequate quantity. It is possible that socially disadvantaged women discover early in life the low quality of the health services provided for them and decide that the costs of making visits (e.g., child care, transportation, opportunity costs of time spent) exceed the benefits.

From a policy perspective, this finding means that the prenatal care problem in this country rests as much with the quality of care delivered to disadvantaged women in particular as with this group's failure to seek the care that is available. Several approaches to the problem are apparent. We need to ascertain, first, that protocols for prenatal care at public facilities match the ACOG guidelines and, second, that these protocols are being followed by physicians and hospital staff. Third, we need to develop a quality control system for private providers. Federal and state funders of prenatal care should make quality assurance an integral part of their evaluation protocol. Finally, low reimbursement rates for prenatal care may contribute to the differences in surveillance outlined here. Providers of services to Medicaid recipients have an incentive to take shortcuts because of the low payment received for such services. Also, the training of the private providers of prenatal services to the poor may be poorer than that of providers to women with higher reimbursement potential.

On the other hand, women need information about the standard components of prenatal care. Most public education has been directed at the necessity for early prenatal care. The assumption has been that once the woman is integrated into the system, it will provide for her needs adequately. It is clear that for many women, this is an inaccurate portrayal of the situation. Basic information about the normal components of prenatal care could be incorporated in the educational campaign that promotes participation in the prenatal care system. Undoubtedly, well-informed and socially advantaged women not only choose better providers, but place demands on providers to deliver more services. Although women, especially socially disadvantaged women, cannot be expected to act as powerful enforcers of medical standards, they may at least be able to change providers if they are dissatisfied. ☐

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Sixth National Conference on Chronic Disease Prevention and Control to Be Held in Washington, DC

The Centers for Disease Control, the Association of State and Territorial Health Officials, and the Association of State and Territorial Chronic Disease Program Directors are the cosponsors of the Sixth National Conference on Chronic Disease Prevention and Control entitled "Making Prevention a Reality." The conference will be held October 22 to 24, 1991, in Washington, DC, and is open to the public

The conference will emphasize interactions between federal, state, and local health departments; voluntary health agencies; and professional organizations. It will include two plenary sessions: Legislative Issues in Chronic Disease Prevention, and Community Action in Minority Communities for Achieving the Year 2000 Objectives. Concurrent sessions of invited papers will focus on the following topics:

- Use of Epidemiology for Program Decision Making
- Women's Health Issues—Cardiovascular Disease, Osteoporosis, Aging
- Breast and Cervical Cancer Control
- Behavioral Research—What Have We Learned?
- Linkages Between Academic Centers and Public Health Agencies
- State Activities in Community Chronic Disease Prevention

Additional information is available from the Center for Chronic Disease Prevention and Health Promotion Mailstop K-43, CDC, Atlanta, GA 30033 (tel: 404/488-5390 or FTS 236-5390; fax: 404/488-5962).

Effects of Paternal Occupational Exposure on Spontaneous Abortions

Marja-Liisa Lindbohm, MSc, Kari Hemminki, MD, Michèle G. Bonhomme, MSPH, Ahti Anttila, MSc, Kaarina Rantala, MSc(Eng), Pirjo Heikkilä, MSc, and Michael J. Rosenberg, MD

ABSTRACT

Background: Paternal exposure to mutagenic agents has been suggested to affect pregnancy outcome adversely.

Methods: A nationwide data base of medically diagnosed spontaneous abortions and other pregnancies and national census data was used to evaluate the effects of men's occupational exposures on risk of spontaneous abortion in 99 186 pregnancies in Finland. Census data from the years 1975 and 1980 provided information about the occupation, industry, and socioeconomic status. A job-exposure classification was developed to classify women and their husbands according to possible occupational exposures on the basis of their occupational title and industry.

Results: In 10% of the pregnancies, the husband was exposed to one or more of the mutagens, and the rate of spontaneous abortion was unaffected (OR = 1.0). Of the 25 specific mutagenic exposures evaluated, paternal exposure to four (ethylene oxide, rubber chemicals, solvents used in refineries, and solvents used in the manufacturing of rubber products) was associated with an increased relative risk of spontaneous abortion. In addition, the risk of spontaneous abortion was higher among wives of rubber products workers than among unexposed men.

Conclusions: Although there is some biological rationale for the findings of this study, these findings need to be confirmed by studies in which individual exposures can be measured directly. (*Am J Public Health*. 1991;81:1029–1033)

Introduction

In animals paternal exposure to mutagenic compounds can increase the rate of spontaneous abortions.[1] In humans, however, this relationship remains unclear. For vinyl chloride,[2] anesthetic gases,[3] dibromochloropropane,[4] chloroprene,[5] smelter work,[6] waste water exposures,[7] and organic solvents,[8] effects have been suggested, but the data either have been contradictory or remain unconfirmed. An association has also been found with some paternal occupations: metal-plate and constructional steel workers, crushers and grinders, sewers, workers caring for fur-bearing animals,[9] and mechanics and repairers of motor vehicles.[10]

Several methodologic issues must be considered in conducting studies concerning male effects on reproduction. First, such studies must take into account maternal factors. Second, they should collect data on pregnancy outcomes which minimize systematic bias. Third, they need large enough numbers of exposed men to provide reasonable statistical power. We attempted to address these issues in studying the effects of male occupational exposures on the outcome of pregnancy by using data from two national censuses and a nationwide data base of medically diagnosed pregnancy outcomes. Exposures were determined indirectly by using a job-exposure classification. This large dataset can be used to generate hypotheses for other studies of the effects of occupational exposures on reproductive outcomes.

Methods

Data on reproductive outcomes and maternal age were obtained from the nationwide Hospital Discharge Register (HDR) and from hospital clinics. We collected data on all women with a diagnosis of spontaneous abortion (ICD-8 codes 643 and 645), induced abortion (ICD-8 codes 640–642) and birth (ICD-8 codes 650–662) between 1973 and 1982. A separate questionnaire was sent to the hospitals to collect information on patients treated on an outpatient basis and not registered in the HDR. The content, technical quality, and reliability of the data base of medically diagnosed pregnancies have been described elsewhere.[11]

Information on the occupation and industry of women and their husbands and the women's socioeconomic status was obtained from the 1975 and 1980 national censuses of the Central Statistical Office of Finland. Pregnancy data for the years 1973 to 1978 were linked to the 1975 census, and the data for the years 1979 to 1982 to the 1980 census.

In the census of 1975, performed on December 31, 1975, occupational information was requested as follows: "Occupation in the place of employment" with the instruction "State your occupation as accurately as possible, not only by hon-

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Marja-Liisa Lindbohm, Kari Hemminki, Ahti Anttila, Kaarina Rantala, and Pirjo Heikkilä are with the Institute of Occupational Health in Helsinki, Finland; Michèle G. Bonhomme is with Family Health International, Research Triangle, North Carolina, and Michael J. Rosenberg is with Health Decisions, Inc., Chapel Hill, North Carolina.

Requests for reprints should be sent to Marja-Liisa Lindbohm, Institute of Occupational Health, Topeliuksenkatu 41 a A, SF-00250, Helsinki, Finland.

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TABLE 1—Pregnancy Outcomes Among Wives and Adjusted[a] Odds Ratio of Spontaneous Abortion for Paternal Exposure to Mutagenic Agents

Husband's Exposure Status	Spontaneous Abortions	Induced Abortions	Births	All Pregnancies	Odds Ratio	95% Confidence Interval
Not exposed[b]	7772	7029	72 8[illegible]	87 616	1.0	
Potential, low exposure	820	1045	806[illegible]	9930	1.0	0.9 – 1.0
Moderate or high exposure	139	151	1350	1640	1.0	0.9 – 1.2

[a] Adjusted for age (≤24, 25–29, 30–34, 35–39, ≥40); socioeconomic status (1 = employers, own-account workers, upper level employees, 2 = lower level employees, students and pupils, 3 = manual workers); maternal exposure to reproductive hazards (not exposed, exposed).

[b] Referent category.

orary titles or degrees." In the census of 1980, performed on November 1, 1980, the question was, "In which occupation did you work for the longest period during the last 12 months?" Occupations were coded by census officials, on the basis of the Nordic Classification of Occupations and the International Standard Classification of Occupations. Industry was coded on the basis of the Finnish Standard Industrial Classification.

We classified the occupational exposure status of men and women on the basis of occupation and industry, using a job-exposure classification developed in cooperation with two experienced industrial hygienists. The job-exposure classification is organized by job groups, which are combinations of occupation and industry having a similar type of exposure. For each job group a list is given of agents suspected of having adverse effects on reproduction. The assessment of the exposures among the job groups was based on the industrial hygienic measurements made by the Institute of Occupational Health[12,13] and the Finnish register of employees occupationally exposed to carcinogens (ASA Register).[14]

The list of potentially dangerous agents in the classification comprises 78 exposures, including separate chemical substances (e.g., benzene, ethylene oxide), mixtures of exposures (e.g., mixtures of solvents or metals), and nonspecific exposure groups (e.g., biologic factors, rubber chemicals, pesticides). Twenty-five of the 78 exposures were suspected mutagens, based mainly on the classification of the International Agency for Research on Cancer.[15]

Three levels of exposure were distinguished according to the results of industrial hygienic measurements and the prevalence of exposure in the job group: "moderate or high," "potential, low," and "none." "Moderate or high exposure" included jobs in which the level of exposure to mutagens was continuously at least half of the Threshold Limit Value (TLV) or higher or in which the exposure periodically exceeded TLV and the prevalence of exposure was high. "Potential, low exposure" denoted either (a) jobs with low level but high prevalence of exposure to mutagens, (b) jobs which lacked industrial hygiene measurements but which were reported to the ASA Register, or (c) jobs with a high level and unknown prevalence of exposure.

The susceptible time period was determined on the basis of the time of conception and the time period for which the two censuses obtained information on occupation. For men, the time of spermatogenesis (80 days before conception) was defined as the relevant time period for exposure; for women, the relevant time period was considered to be the first trimester of pregnancy. The analysis of exposure to any mutagenic agents was limited to the pregnancies terminated between January 1, 1976, and December 31, 1976, and between May 1, 1980, and April 30, 1981. In the analysis of specific mutagenic exposures we used a 2-year period close to the census, i.e., January 1, 1976, to December 31, 1977, and May 1, 1980, to April 30, 1982, because the number of workers exposed to a specific agent was often small.

After the pregnancies of women less than 12 and more than 50 years of age had been excluded as well as those with unspecific data on men's and women's occupation, industry, or socioeconomic status, the study population included in the analysis for the limited time periods of January 1 to December 31, 1976, and May 1, 1980 to April 30, 1981, comprised 99 186 pregnancies.

The risk odds ratios comparing the odds of spontaneous abortion for paternal exposure to a mutagen with the odds of spontaneous abortion for the nonexposed group were obtained by using linear logistic regression analysis.[16] To minimize random error in the effect estimates, low and moderate or high levels of exposure were combined if the number of the workers in either category was too small. Age, socioeconomic status, and maternal exposure to potential reproductive hazards were included in the models to control for the confounding effect of these variables. To gain precision, odds ratios were only adjusted for age if fewer than 100 pregnancies occurred among the workers exposed to a chemical. The exclusion of socioeconomic status in the model did not significantly change the odds ratios for these exposures. Maximum likelihood estimates of odds ratios, which were used as surrogate measures of relative risks, and their 95% confidence intervals were calculated by using the GLIM computer program.[17] Odds ratios for exposure to mutagens were similar regardless of whether induced abortions were included or excluded from the denominator. In calculating the rate of spontaneous abortion, we included births, spontaneous abortions, and induced abortions in the denominator. (Although it is common in the literature to refer to these measures as rates, they are in fact proportions or risks because they are calculated by dividing the number of spontaneous abortions by the total number of pregnancies.[16])

Results

The rate of spontaneous abortion in the study population was 8.8%, which is nearly the same as the rate among all Finnish women (8.9%).[11] Of the husbands, 10% were considered to have low or potential occupational exposure to mutagenic agents and only 1.7% were classified as heavily exposed (Table 1). After adjusting for maternal age, socioeconomic status, and maternal exposure to reproductive hazards, the odds ratio of spontaneous abortion did not differ significantly from one for either level of paternal exposure to mutagens (Table 1).

The adjusted odds ratio was significantly increased for four paternal exposures: ethylene oxide (OR = 4.7, 95% CI 1.2-18.4), solvents used in the petroleum refineries (OR = 2.2, 95% CI 1.3-3.8), rubber chemicals (OR = 1.5, 95% CI 1.1-2.2) and a subgroup of the former, solvents used in the manufacture of rubber products (OR = 1.9, 95% CI 1.2-2.8; Table 2). Except for ethylene oxide, these exposures are mixtures of several chemicals. From the solvents used in the refineries, gasoline and benzene are suspected to be mutagens. Similarly, from the solvents used in the manufacture of rubber products 1,1,1-trichloroethane and methylene chloride are suspected mutagens. Some of these chemicals also appear in the classification as specific exposures. When paternal exposure to gasoline, benzene, and 1,1,1-trichloroethane were analyzed separately, the odds ratios did not differ significantly from unity.

TABLE 2—Adjusted[a] Odds Ratio of Spontaneous Abortion for Paternal Exposure to Specific Mutagens[b]

Exposure[c] (Level of Exposure)	Number of Spontaneous Abortions	Number of Pregnancies	Odds Ratio	95% Confidence Interval
Ethylene oxide (low)	3	10	4.7[*d]	1.2–18.4
Solvents, petroleum refineries (gasoline, benzene) (moderate or high)	16	93	2.2[**d]	1.3–3.8
Impregnents of wood (arsenic, chromium, creosotes) (low and moderate or high)	5	33	1.9[d]	0.7–5.0
Solvents, manufacture of rubber products (1,1,1-trichloroethane, methylene chloride) (low and moderate or high)	26	180	1.9[**]	1.2–2.8
Rubber chemicals (low and moderate or high)	34	280	1.5[*]	1.1–2.2
1,1,1-Trichloroethane (moderate or high)	3	24	1.5[d]	0.4–5.0
Radon (moderate or high)	26	248	1.2	0.8–1.8
Formaldehyde (formaline)				
Low exposure	110	1212	1.1	0.9–1.4
Moderate or high exposure	54	596	1.0	0.8–1.4
Polycyclic aromatic hydrocarbons				
Low exposure	1022	11 970	1.0	0.9–1.0
Moderate or high exposure	142	1619	1.0	0.9–1.2
Nickel and nickel oxides (low)	284	3371	1.0	0.9–1.1
Chromium and chromium compounds (low)	286	3444	1.0	0.9–1.1
Solvents, manufacture of plastic products (styrene) (moderate or high)	24	285	1.0	0.7–1.5
Benzene (low)	55	655	1.0	0.7–1.3
Metals (nickel, chromium, cadmium), basic metal industries and manufacture of metal products (low)	155	1603	1.0	0.8–1.1
Motor gasoline (low)	189	2381	0.9	0.8–1.1
Styrene (moderate or high)	9	120	0.9	0.5–1.8
Trichloroethylene (low and moderate or high)	5	66	0.9[d]	0.3–2.1
Plastic monomers (low)	10	158	0.8	0.4–1.4
Tetrachloroethylene (moderate or high)	3	45	0.7[d]	0.2–2.4
Solvents, manufacture of drugs (chloroform, methylene chloride) (moderate or high)	3	60	0.5[d]	0.2–1.7
Raw materials of glass (arsenic, cadmium) (low)	4	99	0.5[d]	0.2–1.3
Hydrazine (low)	2	51	0.4[d]	0.1–1.7
Solvents, manufacture of printing inks (ethene) (moderate or high)	1	33	0.3[d]	0.0–2.5
Nitroglycol, nitroglyserine, trinitrotoluene (low)	1	34	0.3[d]	0.0–2.3
Solvents, manufacture of glue and gasein (ethene) (moderate or high)	1	35	0.3[d]	0.0–2.3

[a] Adjusted for age (≤24, 25–34, ≥35); socioeconomic status (1 = employers, own-account workers, upper level employees, 2 = lower level employees, students and pupils, 3 = manual workers).

[b] Comparison group was not exposed to any mutagen.

[c] In parentheses are the exposures suspected to be mutagens.

[d] Adjusted for age only.

*$P < .05$, **$P < .01$

The risks of spontaneous abortions were also calculated for paternal exposure to lead, carbon disulfide, and heat, which are not mutagens but may adversely affect spermatogenesis. For all these exposures, the odds ratio was below unity but the corresponding confidence interval did not exclude one (Table 3).

In some cases, job title may be a good and specific index of a worker's exposure. For this reason, we calculated the odds ratios of spontaneous abortion by paternal occupation using the 55 occupations included in the exposure classification. The comparison group consisted of wives whose husbands were not exposed to any mutagen. Table 4 lists in descending order the paternal occupations with the highest odds ratios (> 1.20).

Only the results obtained for wives of men employed as rubber product workers reached statistical significance. At the top of the list are also two other occupations in which workers may be exposed to chemicals which we found were associated with an increased risk of abortions (Table 2). These occupations are technical nursing assistants (potentially exposed to ethylene oxide) and oil refinery workers (potentially exposed to solvents of oil manufacture). Other occupations with a high risk of abortion were watchmakers and chimney sweeps.

Discussion

Paternal exposure to certain mutagens may induce spontaneous abortions. We found no difference in the risk of spontaneous abortion between wives of men exposed to any mutagen and wives of men exposed to no mutagen. This may suggest that only a subgroup of mutagens induce spontaneous abortions. Moreover, the mutagenicity of some of the classified agents, such as 1,1,1-trichloroethane, trichloroethylene, tetrachloroethylene, chloroform, methylene chloride, and the

TABLE 3—Adjusted[a] Odds Ratio of Spontaneous Abortions for Exposure to Lead, Carbon Disulfide, and Heat[b]

Exposure (Level of Exposure)	Number of Spontaneous Abortions	Number of Pregnancies	Odds Ratio	95% Confidence Interval
Lead (low and high)	352	4476	0.9	0.8–1.0
Carbon disulfide (high)	2	30	0.8[c]	0.2–3.3
Heat (low)	4	109	0.4[c]	0.2–1.2

[a] Adjusted for age (≤24, 25–34, ≥35); socioeconomic status (1 = employers, own-account workers, upper level employees, 2 = lower level employees, students and pupils, 3 = manual workers).
[b] Comparison group was not exposed to any mutagen, lead, carbon disulfide, or heat.
[c] Adjusted for age only.

TABLE 4—Adjusted[a] Odds Ratio of Spontaneous Abortions in the Occupations of the Exposure Classification with Odds Ratio > 1.20[b]

Occupation	Spontaneous Abortions	Pregnancies	Odds Ratio	95% Confidence Interval
Technical nursing assistants	3	22	1.6	0.5–5.6
Rubber product workers	35	281	1.5[*]	1.1–2.2
Watchmakers	13	102	1.5	0.8–2.7
Chimney sweeps	18	141	1.5	0.9–2.4
Refinery workers	13	111	1.4	0.8–2.5
Smelters, hardeners, temperers in metal industry	4	37	1.4	0.5–3.9
Textile finishers and dyers	14	130	1.3	0.8–2.4
Wire and pipe drawers	5	50	1.2	0.5–3.1
Textile machine setters, operators	22	217	1.2	0.8–1.9

[a] Adjusted for age (≤24, 25–34, ≥35 years).
[b] Comparison group was not exposed to any mutagen.
[*] $P < .05$.

metals as a group is not well established and depends on the mutation assay used. Thus mutagenicity of the agents such as these is debatable.[15,18] Another example of lack of effect is paternal smoking, one of the most common mutagenic exposures, which appears to have no effect on fertility (time to pregnancy).[19,20]

Of the 25 specific mutagens or groups of mutagens evaluated in this study, paternal exposure to the following four was associated with an increased risk of spontaneous abortion: ethylene oxide, rubber chemicals, solvents used in petroleum refineries, and solvents used in the manufacture of rubber products. Among these, ethylene oxide has been identified as a mutagen by almost all mutation assays, including the dominant lethal assay.[15] Exposure to this chemical has also been associated with spontaneous abortion in women who use the chemical to sterilize hospital instruments.[21] However, our finding of the effects of ethylene oxide is based on a small group of exposed workers. In addition, the job group to which exposure to ethylene oxide was attributed is heterogeneous, and no individual exposures could be confirmed. Rubber chemicals contain several microbial mutagens,[22] and an increased risk of abortion has been observed among women exposed to rubber chemicals.[23] An excessive rate of spontaneous abortion has been found among the wives of workers in a waste water treatment plant of a petroleum refinery[7] and among the wives of workers exposed to organic solvents.[8]

Analyses by occupational title identified only 1 out of 55 occupations, rubber workers, for which the wife's risk of spontaneous abortions was higher than that of wives of men not exposed to mutagens. The wives of men exposed to potentially spermatotoxic but nonmutagenic agents, such as lead, carbon disulfide and heat, did not appear to have an increased risk of spontaneous abortions.

Genetic damage to male germ cells is the only mechanism of male-mediated effects on pregnancy outcome for which supporting evidence is available.[24] Another potential mechanism could be that harmful substances are transmitted to the pregnant woman by contact with clothes or by semen leading to secondary maternal exposure. We were unable to separate the effects of these two routes of exposure.

The misclassification of exposure and pregnancy outcome is an important source of bias in studies of occupational exposure and reproductive outcomes. In this study, misclassification of pregnancy outcome was likely to be small because the data were obtained from hospital records, and spontaneous abortions were medically diagnosed. These data were therefore less susceptible to the information bias which might arise from self-reported data on pregnancy outcome.[11]

In Finland induced abortions are legal. If any induced abortions are reported as spontaneous abortions, we would expect this misclassification to be very small.

It is certain that early spontaneous abortions and a small percentage of induced abortions were not included. The duration of pregnancy before a spontaneous abortion may vary by paternal exposure status. If early fetal losses, which were not included in the study, occurred more often among wives of men exposed to mutagens, the odds ratios reported would underestimate the true relative risk. If the converse were true, the odds ratios would overestimate the true relative risk.

The weaknesses of the study pertain to the potential misclassification of exposure status. Assignment of exposure status was made on the basis of occupational titles and industry. In some cases, these data may not have been sufficiently specific to assess exposure status accurately. Consequently, some exposed persons may have been classified incorrectly as unexposed, and this could have resulted in low sensitivity of the exposure assessment. Conversely, occupational categories may have been so broad that unexposed persons may have been classified as exposed, inducing low specificity. Random error in the classification of exposure status (nondifferential misclassification) may have diluted the effects of exposures and may explain the absence of associations in this study.

Although census data do not provide completely accurate exposure information, care was taken to assess exposure as accurately as possible. First, analyses were carried out in the time period close to the date of the censuses to allow minimal time for change in occupation. Second, exposure status was determined independently of reproductive outcome. Third, the job exposure classification rather than mere job title was used to assign exposure status; in some cases it was possible to quantify the exposure.

No information was available on other potential confounding factors such as previous spontaneous abortions or the use of alcohol or smoking. Probably the adjustment of the socioeconomic status has partly controlled the effect of smoking and alcohol drinking, as they vary by social class.

This large study showed no overall excess of spontaneous abortions in women whose husbands were exposed to mutagens in general compared with those whose husbands were not exposed to mutagens. However, excesses were observed among wives of men employed as rubber products workers and of men exposed to ethylene oxide, solvents used in petroleum refineries, rubber chemicals, and solvents used in the manufacturing of rubber products. Many comparisons were made in this study, and some of the associations probably reflect chance alone. Thus, our findings are only suggestive. This emphasizes the need to base inferences on biologic plausibility and on the evidence accumulated from this and other studies. Although there is some biological rationale for these findings, they need to be confirmed by studies in which individual exposures can be assessed directly. ☐

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The Health Hazards of Saunas and Spas and How to Minimize Them

Edward Press, MD, MPH

ABSTRACT

Background: The rapidly increasing number of spas, hot tubs, and saunas intensifies the potentials for deaths from hyperthermia and drowning.

Methods: I analyzed 54 such deaths reported to me by 55 medical examiners and coroners in the United States and 104 deaths reported to the US Consumer Product Safety Commission (CPSC).

Results: Only seven of the 158 deaths analyzed occurred in saunas. All of the remaining deaths occurred in spas, jacuzzis, or hot tubs, which were far more numerous. The chief risk factors identified were alcohol ingestion, heart disease, seizure disorders, and cocaine ingestion (alone or in combination with alcohol ingestion). These factors accounted for 71 or 44.7% of the 159 fatalities. Of these risk factors, alcohol represented 38%; heart disease, 31%; seizure disorders, 17%; and cocaine ingestion, alone or in combination with alcohol, 14%. Sixty-one of the 151 spa-associated deaths occurred in children under 12 years of age. Accidental drownings from uncovered or improperly covered spas and,to a lesser extent, entrapment by suction, were the chief causes of childhood drownings.

Conclusions: Children and older persons who have heart disease or seizure disorders or who use alcohol or cocaine are especially vulnerable. Recommended preventive measures include shortening the time of exposure, lowering the temperature, establishing safety standards for covers and for baffles for suction outlets, and using warning notices. (*Am J Public Health.* 1991;81:1034–1037)

Introduction

The "dry" or Finnish Sauna is an enclosure large enough for two to eight or more persons; it generally has no active ventilation, is heated by electricity, gas, or other means; temperatures are maintained at 70°C to 90°C, but often can be raised up to 100°C. Spas are large bath tubs or small pools made of synthetic materials; they have equipment that heats, recirculates, and disinfects the water and usually has jets of water or air or both. Hot tubs are similar to spas, but they are made of wood and their recirculating water may or may not have jets of water or air. The water in both spas and hot tubs is heated to about 40°C, but may go even higher. Health hazards from hot tubs are similar to those from spas. In this article the term "spas" includes hot tubs, but does not include the smaller jetted bath tubs that are emptied after each use.

The exact number of saunas in the United States is not known. The Sauna Society of America estimates the number at about 1 million (personal communication from V.S. Choslowsky, President, Sauna Society of America, Sept. 1988). More accurate estimates are available for spas and hot tubs; one of the trade associations placed the number at about 1.8 million in 1988 (personal communication from Larry E. Paulick, Sr., vice-president, National Spa and Pool Institute, Alexandria, Va, Dec. 5, 1988). A trade publication placed the number that are sold annually at 250 000 (personal communication from Kathleen Bedenbaugh, editor, Aqua Magazine, Harcourt-Brace Jovanovich Publications, Santa Anna, Ca, Oct. 1988. Thus there would be millions of monthly exposures to the hazards of hyperthermia and drowning. To reduce the risks from these exposures, voluntary standards have been developed by several organizations. One of these, Underwriters Laboratories' Standard 875 for Electric Saunas requires the use of the following warning sign: "Do not exceed 30 minutes in sauna. Persons with poor health should consult physician before using sauna." Because (as noted below) a 30-minute exposure has been found to be unsafe for some persons, and because there is so little information available on mortality in saunas and spas in the United States, I undertook a literature review and then attempted to collect information on deaths associated with spas and saunas.

Literature Review

According to Kauppinen and Vuori,[1] warmth induces a feeling of euphoria, relaxation, and tranquility in many persons. Sorri discussed saunas in a regular issue of the Annals of Clinical Research[2] and in a special issue of the same journal,[3] which includes two editorials and 16 articles on the Finnish sauna. The author states that bathing in a sauna is a pleasant and relaxing experience, which combines psychic, physical, and social pleasures; reduces aggressive behavior; and enables bathers to forget the common pressures of everyday life. He also feels that the nature of the experience makes it difficult to define objective scientific methods of measurement.

Edward Press is Professor Emeritus of Public Health and Preventive Medicine at the Oregon Health Sciences University, School of Medicine.

Requests for reprints should be sent to Edward Press, MD, 2211 SW First Avenue, Suite 901, Portland, OR 97201-5013.

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The increased pressure from the water results in peripheral vasodilation, profuse perspiration with a decrease in the extracellular blood volume, an increased viscosity of the blood, an increase in hematocrit with shortened clotting time, and a 20% increase in the consumption of oxygen.[1] The electrocardiographic (ECG) changes include a T-wave depression, ventricular and supraventricular extrasystoles, and a deep S wave.[4] Thus the possibility of cardiac arrest exists. This would be especially true for persons with preexisting heart disease or those on medications that influence cardiac rhythms. The combination of peripheral vasodilation and decreased blood volume from the loss of fluids reduces the cerebral blood flow and increases the likelihood of syncope. Three out of 60 subjects using a Finnish Sauna for 20 minutes at a temperature of 80°C to 90°C fainted; a fourth, a 63-year-old, apparently healthy, man, developed an angina attack; his ECG tracings showed ventricular and extraventricular extrasystoles, which persisted for a half hour after he left the sauna.[4] Jokinen et al.[5] studying children found that of 61 children that had a 10-minute exposure in a sauna, two developed vasovagal collapses.

Diabetics receiving subcutaneous injections of insulin may absorb their insulin much more rapidly because of the peripheral vasodilation. Khogali and Gales[6] feel that "it is quite common for patients with heat stroke to have diabetes, as though they are predisposed. . . . Since diabetics often exhibit postural hypotension due to faulty baroreflexes it is conceivable that this could be responsible for the relatively poor adjustments in vasculature of various body regions." Sweating abnormalities and neuropathy of the autonomic nervous system also are common to diabetics. Hormonal activity resulting in increased secretions of catecholamines and an increased sympathomimetic activity have been reported.

Because in the sauna and spa the testicles, which are normally below internal body temperature, are exposed to temperatures well above normal body temperature, the possibility of impaired spermatogenesis exists. B. J. Procope, as quoted by Kauppinen and Vuori[1] stated that intensive sauna bathing, repeated six to eight times within a 2-week period, caused a reduction in the number of sperm cells in 12 medical students accustomed to the sauna. Brown-Woodman, also quoted by Kauppinen and Vuori, reported that even a single 20-minute exposure to a sauna at 80°C decreased the number of sperm cells starting 1 week after the bathing and lasting for 5 weeks. He found that there were no changes in the morphology; but ultrastructural changes of the sperm cell were observed.

Questions of potential teratologic effects on the developing fetus when pregnant women are exposed to the hyperthermia of the sauna and the possibility that such exposure might induce labor prematurely in women approaching full term, have also been raised. Edwards[7] in a fairly extensive review of the clinical significance of experimental studies on hyperthermia as a teratogen, found that "the species known to be affected include birds, all the common laboratory animals, farm animals, and primates. It would be remarkable if it were not also a human teratogen." He concluded that because there is very strong evidence that hyperthermia is a human teratogen more study is required to fully determine the nature and extent of this risk to prenatal development. He thought that an elevation of 1.5°C to 2.5°C above the species' normal temperature could result in a heat-induced teratogenic effect.

Vaha-Eskeli and Erkkola[8] after reviewing some of the literature and testing 23 pregnant women, concluded that at least, as practiced in Finland, hyperthermia was not teratogenic. The authors did, however, comment that in Finland it is commonly believed that sauna bathing could be used to induce labor in postterm women. In their study of the effect of thermal stress on 23 women who were 36 weeks pregnant, they found that 5 of the women had occasional uterine contractions, although 1 of the 5 experienced the contractions before, as well as after, the sauna exposure.

Method

After a preliminary telephone poll of 16 US medical examiners and coroners, a questionnaire (available from the author) requesting information on the number and type of deaths associated with saunas, spas, and hot tubs and related risk factors noted during the past 5 years was sent to 184 members of the National Association of Medical Examiners. The members were selected on a subjective basis from the entire list of 654 members. When several different members were from the same office, the chief or director, rather than deputies, assistants, or consulting pathologists was chosen. In addition, wide geographic distribution coupled with representation from urban and rural areas and from hot and cold climates were factors in the selection. To supplement these factors, a medical examiner who was a longtime member of the association and familiar with many of its members personally chose about 10 or 12 whom he considered experienced and well trained.

A total of 55 responses were received. These were followed up by telephone interviews with 20 of the 55 for more complete and detailed information. Thirty-two of the 55 respondents indicated that they could remember no fatalities during the preceding 5 years (from November 1988).

The Consumer Product Safety Commission (CPSC) has a program called the Medical Examiners and Coroners Alert Program asking medical examiners and coroners to report deaths to the CPSC. Reports of such deaths in saunas and spas were obtained from the CPSC for the period of January 1, 1986 to November 4, 1988. The reports did not include all states in the United States for the entire period, nor did those reporting necessarily include all cases in their jurisdiction.

Results

Age distribution. Only 15% of the deaths reported in my survey (all in spas), were in children under 12 years of age, compared with 52% reported to the CPSC (Table 1). This discrepancy could reflect the fact that the reports to the CPSC were more recent, most of them prepared in 1987 and 1988 when a larger number of spas were sold to private residences than to public facilities and athletic clubs, which are much less likely to have children present. The author's poll went back 5 years from 1988 when the number of private spas was lower.

Duration of exposure. Ten of the 24 jurisdictions reporting fatalities in my survey recorded the length of exposure to the spa prior to death. In four cases the exposure was more than 30 minutes; in four more, it was 10 to 20 minutes; and in two it was less than 10 minutes.

Risk Factors. Risk factors were reported in 45% of the deaths (Table 2). Alcohol ingestion and heart disease were the chief factors associated with the fatalities. Seizure disorders and cocaine ingestion either alone, or cocaine combined with alcohol, accounted for additional risk factors. In many deaths the presence or absence of risk factors was either unknown or unrecorded.

TABLE 1—Age Distribution of Spa and Sauna Fatalities

Age	SPAS CPSC[a] N	SPAS CPSC[a] %	SPAS EP Survey[b] N	SPAS EP Survey[b] %	SAUNAS EP Survey[b] N	SAUNAS EP Survey[b] %
0–2 y	36	35%	2	4%	0	0%
2–11 y	18	17%	5	11%	0	0%
12–65 y	31	30%	29	62%	2	29%
over 65 y	19	18%	9	19%	5	71%
Unknown	0	0%	2	4%	0	0%
Total	104	100%	47	100%	7	100%

[a] US Consumer Product Safety Commission reports, Jan 1, 1986, to Nov 4, 1988.
[b] Reports to the author from 55 county and state medical examiners and coroners, Nov 1988.

TABLE 2—Risk Factors Reported as Associated with Spa and Sauna Deaths

	SPAS CPSC[a]	SPAS[b]	SAUNAS[b]	TOTAL
Alcohol	10	15	2	27
Heart Disease	10	8	4	22
Seizures	6	6		12
Cocaine	1	2		3
Cocaine & alcohol	3	4		7
Total Reported	30	35	6	71 (44.65%)[c]

[a] Consumer Product Safety Commission's death certificate reports from their Medical Examiners and Coroners Alert Program. In only 50 of the 104 deaths were the risk factors checked as present or unknown. In the others, no mention was made of them.

[b] Author's poll of 55 medical examiners or coroners. In this poll, obesity was recorded in association with heart disease or alcohol in 8 cases.

[c] $\frac{71}{104+55} \times 100 = 44.65\%$

Discussion

The chief physiological differences between spas and saunas relate to the more rapid transfer of heat from water than from air. The jet currents in the spa further enhance the heat transfer and the evaporative cooling effects of perspiration that occur in the dry sauna do not occur when most of the body is under water in spas. Although the transfer of heat is more rapid in spas than in saunas, the gradient of heat is less. Water in spas is generally not above 40°C, which is in contrast to the 80°C or 90°C temperature of the Finnish sauna or the 50°C temperature of the steam room or Japanese sauna.

In addition to the problems of hyperthermia the hazard of drowning exists in spas. Infants can accidentally fall into the spa and drown when the spa is not adequately covered; or older children can become entrapped if the powerful outflow suction is not properly shielded. As noted in Table 2, over half of the 104 deaths reported to the CPSC were in children under 12 years of age. Shinaberger et al.[9] in a study of 74 deaths of young children in California spas found that the major factors involved were lack of supervision, access to the spas, neuromotor handicaps, and entrapment by suction.

Syncope is a hazard for adults. Hyperthermia may induce syncope, particularly in persons with heart disease or alcohol or drug ingestion. Even without these factors, staying in the spa too long or allowing the temperature to rise beyond safe limits can cause syncope.

The greater likelihood of narrowed arteriosclerotic blood vessels in the heart and brain of the elderly poses a risk, especially if they are perspiring profusely and, as a result, become dehydrated and have increased blood viscosity. Alcohol or cocaine, medications that affect the heart or the temperature-regulating mechanism or that result in drowsiness are more likely to be present in adults. Trying to sober up or counter a hangover by "sweating it out" can be a perilous activity. Ylikhari et al.[10] noted that in Finland, which has a population of 4.8 million, the consumption of alcohol has been estimated to be a contributing factor in 20 to 25 sauna-related deaths each year. He felt that alcohol intoxication and particularly the hangover phase exposes a person to cardiac arrhythmias and that the sauna may further increase the arrhythmia risk because of enhanced adrenergic activity.

The length of stay in a sauna or spa is important. The heating switch on some electrically heated saunas is set to go off automatically after 30 minutes, and one of the standards (currently under revision) cautions against exceeding 30 minutes. However, Sohar et al.[4] found even 20 minutes to be too long for 4 of his 60 subjects as noted above. The New York City Health Department requires posting a notice in public saunas, steam rooms, and spas with the following warning: "Use . . . should not exceed 15 minutes. Excessive exposure can be harmful to health" (personal communication, November 1987, between the commissioner of the New York City Health Department and the chair of the CPSC). The president of the Institute for Health Promotion Research in Finland who has done research and had an unusually extensive experience with saunas stated that "—even 10 minutes may be too long a sauna exposure for some individuals." In 2 of the 10 fatalities in the author's survey where the time was recorded the time was less than 10 minutes. Studies to determine how long pregnant women can safely remain in a spa found that the length of time varied with the temperature of the water.[11] However, using 38.9°C as a potentially teratogenic temperature, Harvey et al.[11] found that in a hot tub with water at 39°C it took the women 15 to 25 minutes (mean 23.0 minutes) to reach a body temperature of 38.9°C. When the water temperature was 41.1°C it took the women 10 to 30 minutes (mean 18.5 minutes). In a sauna at 81.4°C all 20 women felt sufficient discomfort (dizziness, tingling in hands, rapid pulse, irregular heart beat, and stomach pain) to leave before their temperature rose to 38.9°C. The subjects in the 39°C hot tubs began to leave because of discomfort after 10 minutes; those in the 41.1°C tub left after 5 minutes. Those who elected to remain because they did not experience any discomfort required the above noted times to have their temperatures reach 38.9°C.

Spas, sometimes marketed through discount stores at cut rates are often purchased by families with small children. The combination of inadequately supervised children and the use of cheaper models with substandard covers and poorly shielded suction outlets could result in accidental drowning. Mandatory, rather than voluntary, safety standards for covers and baffles might decrease this hazard. Many of the standards developed by several organizations need to be revised, and for all compliance is now voluntary. Standards have been formulated by Underwriters Laboratories[12] and the National Spa and Pool Institute.[13] The Centers for Disease Control of the US Department of Health and Human Services has also published some recommended health and safety guidelines for public spas and hot tubs.[14] A few state or local jurisdictions have adopted some of these standards (with or without recent revisions) as requirements for public facilities; they may also enforce them. However, most official agencies have taken no action. Moreover, the current mandatory standards of most official agencies cover only public and not private residential facilities. The CPSC has the authority and responsibility to promulgate mandatory standards for both public and private saunas and spas but, as of December 1990, has not done so.

Both the author's survey and the CPSC data suffer from incomplete information: the total population at risk, total number of associated deaths, and the number, type and frequency of use are not known. The sparse response rate to my survey as well as its lack of complete coverage means that many deaths and many risk factors may have been overlooked. A more thorough and complete study is needed, preferably a prospective one, to obtain a more accurate and complete picture of the problem.

Nevertheless, even the limited data presented here suggest that a public health problem exists and that the problem can be corrected through appropriate actions by the CPSC and official health agencies.

Persons with heart disease, hypertension, seizure disorders, diabetes, or significant obesity; persons who have ingested alcohol, narcotic drugs, or medications that can result in drowsiness or interfere with the body's temperature-regulating mechanism; or persons who are over 65 years of age must limit their stays in saunas or spas to 5 or at the most 10 minutes at a time. Similar precautions should be observed by women in the first trimester of pregnancy. Even healthy adults would be well advised not to stay in for more than 10 to 15 minutes at a time.

Automatic controls, limiting air and water temperatures to 80°C for the Finnish Sauna, 50°C for the steam room, and 40°C for the water temperature in the spas should be required as should effective covers for spas and proper baffles or shields for suction outlets. ☐

Acknowledgment

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Serum Total Cholesterol: HDL Cholesterol Ratios in US White and Black Adults by Selected Demographic and Socioeconomic Variables (HANES II)

Shai Linn, MD, DrPH, Robinson Fulwood, MSPH, Margaret Carroll, MSPH, J. Gerald Brook, MD, Clifford Johnson, MSPH, William D. Kalsbeek, PhD, and Basil M. Rifkind, MD

ABSTRACT

Background: **Framingham Study findings suggest that total cholesterol (TC):High density lipoprotein cholesterol (HDL-C) ratio is a useful summary of the joint contribution of TC and HDL-C to coronary heart disease (CHD) risk. Information on the distribution of TC:HDL-C in the US population is limited to selected populations and the relationship of the ratio distribution and its correlates has received little attention.**

Method: **TC/HDL-C ratios were examined in a representative sample of the United States adult population ages 20 to 74 years, between February 1976 and February 1980 during NHANES II, using stratification and multivariate regression analyses.**

Results: **Age-adjusted mean ratios were higher in men compared with women and were higher in Whites compared with Blacks. White men had the highest TC/HDL-C mean ratios. These relationships remained after stratification by age, education, body mass index, alcohol use, cigarette smoking, and physical activity. Using multivariate analyses, the ratios were positively related to BMI, age, and smoking; and negatively related to female sex, alcohol use, being Black, and physical activity.**

Conclusions: **Using a ratio reference point of greater than or equal to 4.5 from the Framingham study, at least an estimated 44 million persons ages 25 to 74 years in the US were found to be at higher risk of developing coronary heart disease. (*Am J Public Health*. 1991;81:1038–1043)**

Introduction

The various cholesterol-carrying lipoproteins are independent predictors of coronary heart disease (CHD) risk. Plasma total cholesterol (TC) and low density lipoprotein cholesterol (LDL-C) levels are positively correlated with CHD, whereas high density lipoprotein cholesterol (HDL-C) levels are inversely related.[1] However, TC or LDL-C and HDL-C are poorly correlated with each other so that knowledge of both may be important to predict CHD risk than does either of them alone.

Various ratios combining TC or LDL-C and HDL-C have been introduced. Framingham Study findings suggest that TC:HDL-C ratio is a useful summary of the joint contribution of TC and HDL-C to CHD risk.[1–3]

Other summary measures, using LDL-C, may need calculations or special laboratory methods. We have chosen to report the TC:HDL-C to allow comparison with the Framingham data, since it is the main source of information on the relationship of a summary measure of TC and HDL-C with CHD mortality. This ratio has the advantage of being available to the physician and is easier to comprehend at the clinical level.

Information on the distribution of TC:HDL-C (or its inverse) in the United States population is limited to selected populations. Both the Framingham Study[1,3] and the Lipid Research Clinics (LRC) Prevalence Study[4] were not planned to be representative of the US population. Furthermore, the relationship of the ratio distribution and its correlates has received little attention.

The second National Health and Nutrition Examination Survey (NHANES II),[5,6] measured TC and HDL-C levels in a probability sample of the civilian, noninstitutionalized population of the United States and thus provides information on the distribution of the TC:HDL-C ratio for categories of selected potential correlates, namely age, sex, race, education, body mass index (BMI), alcohol intake, smoking, and physical activity.

Methods

The NHANES II, conducted between February 1976 and February 1980 by the National Center for Health Statistics (NCHS), was a national cross-sectional probability survey of the civilian noninstitutionalized population of the United States aged 6 months to 74 years including Alaska and Hawaii.[5,6] It consisted of household interviews, as well as physical examination and interviews in mobile examination centers. The household interview collected socioeconomic

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From the Lipid Metabolism-Atherogenesis Branch, National Heart, Lung and Blood Institute, NIH, Bethesda (Linn, Brook, Rifkind); the Nutrition Statistics Branch, National Center for Health Statistics, Hyattsville, MD (Fulwood, Carroll, Johnson); the Collaborative Studies Coordinating Center, Department of Biostatistics, University of North Carolina, Chapel Hill (Kalsbeek); and the Clinical Epidemiology Unit, Rambam Medical Center, Haifa, Israel (Linn).

Address reprint requests to Basil M. Rifkind, MD, Chief, Lipid Metabolism-Atherogenesis Branch, Division of Heart and Vascular Diseases, NHLBI, NIH, Federal Building, Room 401, Bethesda, MD 20892. This paper, submitted to the Journal April 30, 1990, was revised and accepted for publication January 23, 1991.

and demographic information on the sample person and the family and a medical history questionnaire. The medical examination was conducted according to a standardized protocol and included laboratory tests on whole blood and sera. For each person examined, dietary data including alcohol consumption were obtained by means of a 24-hour recall and a three-month food frequency questionnaire administered by trained dietary interviewers.

The three sources of nonresponse and response rates by sex and race are shown in Table 1. An analysis of selected health variables of persons who were examined versus those who were not examined (but interviewed) indicated no substantial bias due to nonresponse.[7] An examination of the characteristics of persons with and without HDL-C determinations such as age, education, BMI, alcohol use, smoking, and physical activity did not detect differences that would suggest a bias in the results. Adjustments have been made according to the usual NCHS methods so that the results reported here reflect a nationally representative sample.[5]

TABLE 1—NHANES Study population: Numbers and response rates by sex and race

		Men		Women	
	Total*	White	Black	White	Black
Total sample	17058	7059	852	8044	1103
Interviewed	15080	6170	766	7146	998
Response rate (%)	88.4	87.4	89.9	88.8	90.5
Examined	11637	4883	607	5418	729
Response rate (%)	77.2	79.1	79.2	75.8	73.0
HDL determinations	9625	4019	462	4563	581
Response rate (%)	82.7	82.3	76.1	84.2	79.7
Overall response Rate (%)	56.4	56.9	54.2	56.7	52.7

*Includes only Whites and Blacks aged 20–74 years.

Determination of Serum Lipids

Chemical analyses were provided by the LRC Laboratory at the George Washington University, and the editing and data processing were provided by the LRC Central Processing Unit, Department of Biostatistics, at the University of North Carolina. The laboratory techniques have been described elsewhere.[5,8]

Definitions

Educational level: Measured by the highest grade attended in a public or a private school.

Body mass index (BMI, Quetelet index): Weight (Kg)/height (m)}**2.

Frequency of alcohol consumption: Participants were asked about the usual consumption of alcoholic beverages during the three months that preceded the interview. Consumption was classified as follows: none (never), less than once/week (seldom), 1 to 6 times/week (weekly), or 1 or more times/day (daily). The data for all types of alcohol—beer, wine, and liquor—were compared. The greatest frequency of reported alcohol use was calculated as a composite measure of habitual (chronic) frequency of alcohol use of any kind.

Smoking: Interviewers were asked: 1) whether they had ever smoked at least 100 cigarettes during their entire life; 2) whether they currently smoked; and 3) the daily number of cigarettes smoked. Individuals who responded positively to questions #1 and #2 were defined as "current smokers." Those who responded positively to question #1 but negatively to question #2 were defined as "past smokers," and those who responded negatively to question #1 were defined as "nonsmokers." For current smokers, the data were dichotomized into those who smoked fewer than 21 cigarettes per day, and those who smoked 21 or more cigarettes per day.

Physical activity: Interviewees were asked: 1) whether they were getting "much exercise," "moderate exercise," or "little or no exercise" for recreation; 2) whether in their usual day, aside from recreation, they were "very active," "moderately active," or "quite inactive." The higher category of activity, at any time was then calculated as a composite of activity of any kind.

Statistical Methods

All statistical analyses incorporated both the weights and the complex sample design of the NHANES II. Age-specific and age-adjusted mean HDL-C levels of men and women by race and by selected correlates were estimated by use of the programs SURREGR[9] and GENCAT.[10,11] The age-adjusted means were standardized by the direct method to the 1980 US census population. The first level of analysis consisted of testing the equality in age-specific mean and the age-adjusted mean ratio by selected correlates for the four sex-race groups. Hypotheses of no difference in mean ratio were tested at the 0.05 level of significance, using a chi-square statistic with one degree of freedom. For this report the use of the word "significant" implies "statistically significant."

Second, Pearson correlation coefficients were estimated to investigate the relationship between the ratio and continuous variables such as age, years of education, and the number of cigarettes currently smoked. Correlation coefficients and standard errors were estimated by use of the program OSIRIS.[12] Since the sampling distribution of the correlation coefficients is not normal, Fisher's Z transformation of the correlation coefficients was used to test the hypothesis of zero correlation.

Third, using the SURREGR program, weighted least squares multiple regression analyses,[9] with ratio as the dependent variable, were applied to control for potential confounding. All correlates were included simultaneously in the equation.

Results

The mean ratios and selected percentiles by age, sex, and race are presented in Table 2. The ratios were directly related to age ($P < 0.001$) with Pearson correlation coefficients of 0.235 for White men, 0.207 for Black men, 0.301 for White women, and 0.247 for Black women. White men had the highest mean ratios, consistently increasing from 4.11 at ages 20 to 24 to a peak of 5.55 at ages 45 to 54, and declining thereafter.

The ratios for White men were consistently and significantly higher than those of White women. The age-adjusted sex difference for Whites was 0.77 (18 percent higher for males). Black men had higher ratios than Black women at ages 25 to 54, but Black women had higher ratios in the youngest age group 20 to 24 and at age 55 and above. Blacks had consistently lower ratios than Whites, for both men and women. All differences for men were significant.

TABLE 2—Ratio of Total Cholesterol to HDL Cholesterol (mg/dl) of Adults Ages 20–74 Years by Sex, Race, and Age, with Means, Age-Adjusted Means, and Selected Percentiles, United States 1976–80

	Examined Persons	Estimated Population in Thousands	Selected Percentiles									
			Mean	5th	10th	15th	25th	50th	75th	85th	90th	95th
MEN												
Whites												
20–74	4,019	55 806	5.06	2.79	3.17	3.43	3.88	4.81	5.96	6.72	7.29	8.13
Age-adjusted mean ratio			5.05									
20–24	494	8 052	4.11	2.39	2.72	2.95	3.25	3.98	4.75	5.31	5.07	6.51
25–34	751	13 884	4.75	2.58	3.05	3.32	3.74	4.51	5.53	6.22	6.72	7.57
35–44	545	9 808	5.31	3.14	3.41	3.62	4.06	4.97	6.26	6.97	7.79	8.21
45–54	506	9 865	5.55	3.14	3.50	3.86	4.35	5.39	6.49	7.24	7.78	8.79
55–64	865	8 642	5.47	3.08	3.47	3.80	4.17	5.19	6.51	7.25	7.82	8.54
65–74	858	5 576	5.30	2.86	3.32	3.55	4.03	5.11	6.24	6.84	7.39	8.25
Blacks												
20–74	462	6 102	4.28	2.24	2.59	2.77	3.08	4.00	5.11	5.69	6.22	7.36
Age-adjusted mean ratio			4.33									
20–24	65	1 043	3.43	+	2.24	2.44	2.64	3.22	4.20	4.41	4.70	+
25–34	112	1 546	4.16	2.37	2.59	2.74	3.06	3.91	5.04	5.64	6.06	7.00
35–44	54	1 112	4.69	+	2.91	3.34	3.53	4.45	5.28	6.07	6.75	+
45–54	47	1 044	4.75	+	+	2.83	3.54	4.17	5.44	6.09	+	+
55–64	92	801	4.50	+	2.79	3.00	3.38	4.28	5.47	6.30	6.69	+
65–74	92	555	4.49	+	2.67	2.84	3.32	4.13	5.18	5.69	6.64	+
WOMEN												
Whites												
20–74	4,563	60 785	4.30	2.42	2.68	2.88	3.18	3.96	5.04	5.79	6.30	7.24
Age-adjusted mean ratio			4.28									
20–24	542	8 408	3.66	2.23	2.43	2.59	2.90	3.42	4.20	4.64	5.14	5.72
25–34	875	14 494	3.90	2.40	2.58	2.74	2.98	3.63	4.51	5.07	5.58	6.61
35–44	626	10 584	4.20	2.45	2.66	2.86	3.19	3.98	4.97	5.56	6.06	6.74
45–54	546	10 369	4.57	2.45	2.80	3.00	3.33	4.35	5.47	6.05	6.60	7.60
55–64	973	9 601	4.81	2.63	2.91	3.12	3.57	4.52	5.67	6.56	7.18	8.17
65–74	1,001	7 329	5.00	2.88	3.16	3.36	3.83	4.75	5.89	6.73	7.16	8.03
Blacks												
20–74	581	7 579	4.06	2.30	2.51	2.69	2.98	3.70	4.70	5.60	6.26	7.05
Age-adjusted mean ratio			4.10									
20–24	77	1 304	3.45	+	2.39	2.57	2.67	3.22	4.13	4.68	4.95	+
25–34	122	1 953	3.76	2.22	2.38	2.52	2.84	3.41	4.47	5.45	5.86	6.47
35–44	84	1 415	4.12	+	2.45	2.69	2.96	3.87	4.68	5.39	6.52	+
45–54	81	1 215	4.35	+	3.00	3.07	3.19	3.96	5.16	6.42	6.64	+
55–64	99	959	4.78	+	2.60	2.96	3.11	4.21	5.95	7.19	7.64	+
65–74	118	733	4.50	2.52	2.97	3.23	3.59	4.32	5.15	5.78	5.81	7.07

[1] Age-adjusted by direct method to the 1980 U.S. Census Population.

+Sample size is inadequate to produce a reliable estimate.

TC:HDL-C by Categories of Correlates

Education: In Whites, there was a slight initial increase in the ratios from the first to the second category of education, and a consistent decline thereafter. However, for Black men there was a decrease in the ratio from the first to the second level of education, and a consistent increase with higher education. There was no clear trend in the ratio for Black women. The correlation coefficients were negative in all four sex-race groups, suggesting an inverse relationship between education and the ratio (data available on request from authors).

BMI: The ratios increased uniformly in all four sex-race groups with increasing BMI (Table 3). The relationships of the ratio with BMI were stronger in women than in men, and in Blacks than in Whites.

Alcohol: The age-adjusted mean of Whites of both sexes and of Black women decreased consistently with increased frequency of alcohol consumption (Table 4). The ratio of Black men was highest for those who seldom consumed alcohol and decreased consistently thereafter. For each of the four sex-race groups except for Black men there were lower ratios with increasing reported quantities of alcohol consumed over the last 24 hours. The Pearson correlation coefficients were stronger in men than in women and in Whites than in Blacks.

Smoking: In all groups with the exception of Black men, nonsmokers had lower ratios than smokers (Table 5). However, Black men who never smoked and past smokers had higher ratios than light smokers. Among the smokers, the heavier the smoking the higher the ratio in all groups with the exception of Black females, where the number of cases was too small to produce a reliable estimate.

Physical activity: White men and women, and Black men and women had the lowest ratios with highest level of reported physical activity (Table 6).

TC:HDL-C Ratio in the Four Sex-Ratio Groups after Stratification for Correlates

Sex and race patterns remained basically similar to the findings in the basic data even after controlling for correlates: Of the four sex-race groups, White men consistently had the highest ratios; and Black women had the lowest ratios after stratification for BMI and alcohol consumption (data available on request from authors).

Multivariate Analyses

In the multivariate model with all correlates (Table 7), BMI, age, and current smoking were positively and independently related to the ratio. Being female, frequent alcohol use, being Black, and physical activity were negatively and independently related to the ratio. In this model, education was not strongly related to the ratio. Similar results were obtained when the natural log of the ratio was used as a dependent variable.

Discussion

The present report provides estimates of the distribution of the TC:HDL-C ratio and the relationship to selected correlates in a representative sample of the United States population. The ratios were found to increase with age in men and women of both races. The ratios were consistently higher for men compared to women for the two races; and Whites had higher ratios compared to Blacks. These relationships remained even after stratification for categories of correlates.

The age and sex trends of the TC:HDL-C ratios were similar to those seen in the Framingham study which, however, found higher values. The socioeconomic status (SES) of the Framingham population is higher than that of the general US population. However, since higher SES was found to be related to lower ratios in the present study, SES does not account for the differences observed in the two studies. Laboratory and other methodological differences in lipids and lipoprotein determinations could account for some of the differences. A bias in the determination of lipid levels in NHANES II seems to be unlikely. All TC and HDL-C determinations were performed using a standardized procedure, without knowing the participants demographic characteristics and the quality of the laboratory determinations were routinely evaluated.

In the univariate analyses the ratio was negatively related to years of education in the four sex-race groups but when other correlates were considered in a multivariate model, the ratio was only weakly positive and weakly related to education.

TABLE 3—Age-Adjusted Mean Ratio of Total Cholesterol to HDL Cholesterol of Adults, Ages 20–74 Years, by Race, Sex, and Quartiles of Body Mass Index (BMI), and Pearson Correlation Coefficient, United States, 1976–80

Men Quartiles of BMI	*White*	*Black*
Less than 22.7	4.45	3.59
22.7–24.9	4.86	4.07
25.0–27.5	5.25	4.65
27.6 or more	5.67	5.10
Pearson Correlation	0.302*	0.358*
Women Quartiles of BMI	*White*	*Black*
Less than 21.2	3.82	3.25
21.2–23.7	4.06	3.52
23.8–27.4	4.37	3.98
27.5 or more	4.99	4.60
Pearson Correlation	0.313*	0.367*

Age-adjusted by direct method to the 1980 US Census population.
Body mass index quartiles were determined from sex-specific BMI distribution.
*P < 0.001

TABLE 4—Age-Adjusted Mean Ratio of Total Cholesterol to HDL Cholesterol of Adults, Ages 20–74 Years, by Race, Sex, and Frequency of Consumption of Alcoholic Beverages, United States, 1976–80

Frequency of Consumption	White Men	Black Men	White Women	Black Women
Never	5.42	4.44	4.51	4.30
Seldom	5.20	4.65	4.32	4.19
Weekly	4.92	4.19	4.05	3.76
Daily	4.74	3.77	3.63	3.53

Age-adjusted by direct method to the 1980 US Census population.
Frequency of consumption of alcoholic beverage was determined by taking the maximum reported consumption of either beer, wine, or liquor from the food frequency questionnaire.

TABLE 5—Age-Adjusted Mean Ratio of Total Cholesterol to HDL Cholesterol of Adults, Ages 20–74 Years, by Race, Sex, and Smoking Status, and Pearson Correlation Coefficient, United States, 1976–80

Smoking Status	White Men	Black Men	White Women	Black Women
Nonsmokers	4.97	4.46	4.19	4.03
Past smokers	4.99	4.51	4.12	4.21
Smokers				
1–20 cigarettes/day	5.05	4.01	4.45	4.17
21 cigarettes or more/day	5.38	4.54	4.53	—
Pearson Correlation	0.141*	—	0.049	—

Age-adjusted by direct method to the 1980 US Census population.
A past-smoker is defined as a person who had smoked at least 100 cigarettes but does not currently smoke; a nonsmoker did not smoke 100 cigarettes in his lifetime and does not currently smoke.
*P < 0.001.
—Sample size is inadequate to produce a reliable estimate.

Some limitations on the use and interpretation of the ratio may be advised. It is important to emphasize that the TC:HDL-C can change as a function of either TC or HDL-C. Furthermore, the relationship to correlates is complicated by the fact that at any given ratio level a number of correlates can influence infinite combinations of TC and HDL-C and be responsible for the final outcome. Some correlates such as age will have their effect via TC since there is no strong relationship between age and HDL-C. Conversely, alcohol consumption, for

TABLE 6—Age-Adjusted Mean Ratio of Total Cholesterol to HDL Cholesterol of Adults, Ages 20–74 Years, by Race, Sex, and Physical Activity, United States, 1976–80

Physical Activity Level	White Men	Black Men	White Women	Black Women
Quite inactive	5.31	4.31	4.46	4.80
Moderately active	5.12	4.49	4.28	4.06
Very active	4.96	4.12	4.23	3.89

Age-adjusted by direct method to the 1980 US Census population.

TABLE 7—Multiple Linear Regression Model with Ratio of Total Cholesterol to HDL-Cholesterol as a Dependent Including all Correlates

Overall R^2: 0.212 Intercept 1.969	B[a]	partial R^2 (%)
BMI	0.089[b]	9.16
Age (years)	0.022[b]	4.16
Sex (women vs men)	−0.773[b]	4.34
Current smoker	0.409[b]	1.23
Alcohol use (never or seldom vs others)	−0.430[b]	1.13
Race (Blacks vs Whites)	−0.575[b]	1.10
Physical activity (very active or moderate vs inactive)	−0.089[c]	0.07
Education (years)	0.002	0.00

[a] Estimated regression coefficient.
[b] $P < 0.001$.
[c] $P < 0.05$.

example, will predominantly affect HDL-C. BMI on the other hand is known to influence both components of the ratio, and indeed proved to be the most important predictor of TC:HDL-C in the multivariate analyses. Several publications have investigated the relationships of TC or HDL-C to demographic and socioeconomic variables in the US.[5,13–15] When assessing the distribution of the ratio and its correlates, these factors must be taken into consideration.

In the Framingham Study, the TC:HDL-C ratio was found to be an excellent predictor of CHD risk. Men and women who did not develop CHD within the years of follow-up had an average ratio of 5.2. For those who did develop CHD over the same period, the average ratio for men was 5.8 and for women 5.3. The American Heart Association guidelines suggest that to consider all those at or above ratios of 4.5 be considered at higher risk of CHD[1,16] as TC:HDL-C ratio of 4.4 corresponds to total cholesterol level of 200 mg/dl and HDL-C of 45 mg/dl. Using this single reference ratio, Castelli[1] estimated that 50 percent of the women and 75 percent of the men in the United States are at risk of CHD. Better estimates may be obtained using the NHANES II data (Table 2). Thus, 50 percent of the White males and at least 25 percent of the Black males ages 25 and above have ratios at or above 4.5 and would be considered at increased risk of CHD. This corresponds to population estimates of 23,877,500 White males and 1,264,500 Black males. Similarly, at least 25 percent of women, of both races, ages 25 and above would be at higher risk, corresponding to population estimates of 17,326,750 White women and 1,568,750 Black women, respectively. Thus, there are about 44 million White and Black persons in the US ages 25 to 74 who may be at higher risk of developing CHD based on a TC:HDL-C ratio of at or above 4.5. However, it must be emphasized that there is a need to describe the shape of the curve of the TC/HDL-C levels and CHD risk before selecting a reference point to obtain a reliable estimate of the population at higher risk.

The use of TC:HDL-C to predict CHD may be an oversimplification. In only one study in the US, the Framingham Study, has the predictive power of the ratio been determined. Furthermore, it is unclear whether the ratio predicts CHD across all levels of TC.[16,17] □

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Late Breaker Session on Injury Control Invites Abstracts

The Injury Control and Emergency Health Services special primary interest group of the American Public Health Association has announced it will again feature a "late breaker" session during the upcoming 119th APHA annual meeting in Atlanta. The session will be held on Tuesday, November 12, 8:30 AM–10:00 AM, and will feature work completed within the last few months—after the deadline for consideration in the regular symposia of the APHA annual meeting.

Abstracts of 250 words or fewer will be accepted by the Injury Control SPIG until September 10, 1991. Please send the abstract, title of the paper, author's name, address, and telephone number to: Richard Waxweiler, Division of Injury Control, Centers for Disease Control, Mail Stop F-36, Atlanta, GA 30333. Tel: 404/488-4695.

Public Health Briefs

Physicians Promoting Bicycle Helmets for Children: A Randomized Trial

Robert Cushman, MD, FRCP, William James, MD, FRCP, and Helen Waclawik, BScN, RN

ABSTRACT

Head injury is the leading cause of death and serious morbidity in bicycle accidents. There is good evidence to recommend helmets, yet few children wear them. We evaluated helmet promotion in a randomized trial targeting children presenting to primary care settings for routine ambulatory care. The intervention consisted of physician counseling and take-home pamphlets. The study involved 339 families, 167 in the intervention group and 172 in the control group. In a follow-up telephone call, 2 to 3 weeks later, only 7.2% of the intervention group had purchased helmets, compared with 7.0% of the control group (χ^2 = 0.0056, P = .94). During the latter half of the study, bicycle safety received considerable media attention in Ottawa, and the provincial medical society sponsored a $5 discount campaign. Therefore both groups were subject to community "co-intervention." Nonetheless, we were surprised that physician counseling made no additional impact. Our results and the success of certain community programs suggest that physicians interested in helmet promotion would do better to participate in the design and implementation of multidisciplinary campaigns. (*Am J Public Health.* 1991;81:1044-1046)

Introduction

Head injuries account for 70% to 80% of deaths and serious morbidity in bicycle-related injuries.[1-7] Children and young adults are at greatest risk, accounting for more than 50% of deaths, and cycling is well recognized as a leading cause of head injury in childhood.[8-13] There is ample evidence to recommend the use of helmets, yet relatively few cyclists wear them.[14-22] Surveys in Arizona and Ontario found helmet use was less than 2% among elementary and secondary school children, compared with rates of 10% to 20% in the adult population.[10,22]

This study was designed to evaluate the effectiveness of promoting the use of bicycle helmets during a routine primary care visit.

Methods

Study Design

We targeted families with one or more cyclists ages 5 to 18, presenting to their pediatrician's or family doctor's office for a routine ambulatory visit. Physicians exercised discretion and excluded families presenting for emergency care or other special circumstances, where helmet promotion would have been inappropriate. All participants received an envelope containing a brief one-page survey and a consent form. Parents were asked to list the cyclists by age and to provide information about bicycles and helmet ownership and use. Randomly flagged envelopes indicated to the physician which families were to receive the intervention. The control group received only the usual care.

The Intervention

The intervention consisted of helmet promotion counseling and take-home pamphlets. During an individualized orientation session, physicians were provided with a small plasticized cue card outlining facts about bike accidents and head injuries (Table 1). Physicians were also encouraged to use the pamphlets as a focus for discussion and to express their personal concerns about the dangers of riding without protective headgear.

Outcome Measure

At 2 to 3 weeks' postintervention, a research assistant made a follow-up phone call to tabulate helmet purchases. Positive respondents were also questioned about cost, brand, and place of purchase. The answers allowed us to assess the validity of the "yes" responses. Because helmet owners often ride without headgear, additional questions addressed helmet use.[23,24]

Statistical Analysis

We estimated a helmet purchase rate of 5% to 10% in the control group, and for our efforts to be worthwhile, sought a rate 2 to 3 times higher in the intervention group. Based on our calculations, using a

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Robert Cushman is an emergency room physician, William James a pediatrician, and Helen Waclawik an emergency room nurse with the Children's Hospital of Eastern Ontario (CHEO). Dr. Cushman is also a family physician at Dalhousie Health and Community Services.

Requests for reprints should be sent to Robert Cushman, MD, Dalhousie Health and Community Services, 755 Somerset Street West, Ottawa, Ontario, Canada K1R 6R1.

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Editor's Note: See related editorial on p. 972.

TABLE 1—Helmet Promotion Cue Card Contents

1. Last summer over 100 children with bike related injuries were admitted to the children's hospital.
2. Every year more than 600 children in the U.S.A. die from bike accidents.
3. Most serious injuries and deaths result from head injury.
4. Bike accidents are the leading cause of head injury in childhood.

Bike helmets can reduce the number and severity of head injuries. We consider helmets to be standard equipment and we are encouraging all our patients who cycle to purchase and use them.

one-tailed α of .05 and a β of .1 to .2, we determined that a sample size of 300 would provide adequate power. It would allow us to detect an absolute increase of 10% in the helmet purchase rate, from 5% to 15% at the low end of the scale and from 10% to 20% at the high end. The χ^2 was used to compare the number of helmets purchased in the two groups. Baseline characteristics were assessed using χ^2 or t tests where appropriate.

Results

There were 167 families in the intervention group and 172 in the control group. A total of 288 children were allotted to each group. Small differences between the groups were not statistically significant (Table 2).

At follow-up, only 12 (7.2%) of the 167 families in the intervention group had purchased helmets, compared with 12 (7.0%) of the 172 families in the control group. The 0.2% difference in favor of the intervention group was neither clinically nor statistically significant ($\chi^2 = 0.0056$, $P = .94$). Calculation of the 95% confidence interval on this difference gives a margin of error of ±5.5%, meaning that there is only a 2.5% chance that the intervention might have increased the purchase of helmets by 5.7% (0.2 ± 5.5) (Table 3).

Discussion

While the intervention proved disappointing, we were a little surprised by the control group results. During the latter half of the study (summer 1989) helmet promotion received considerable media attention in Ottawa. The Ontario Medical Association, in conjunction with a local department store, offered a $5 discount on helmets with coupons displayed on posters in waiting rooms. Therefore, community "co-intervention" may have influenced the study. We also considered possible "contamination," whereby clinicians inadvertently offered the intervention to the control group. Physicians were briefed about randomization and the two arms of the study. Only the flagged envelopes contained the pamphlets, serving as the cue for counseling, and when we addressed this issue upon completion of the study, participants did not feel contamination was a problem.

The more important question is why helmet promotion produced no additional effect in the intervention group. We considered three possible problems: low-quality intervention, high rates of helmet use, and price. The intervention was deliberately designed for busy office practice. Counseling sessions were brief, and the material provided a framework for physicians to deliver the message in their own style. Physicians were not "hand-picked" for interest in the subject, nor for counseling skills. However, they were volunteers who received a personalized formal orientation, as well as informal monitoring, from the research assistant.

Approximately 26% of families owned at least one helmet at study entry.

TABLE 3—Helmet Purchases by Family

Group	Purchased Helmet	Did Not Purchase Helmet
Intervention	12	155
Control	12	160

On a per capita basis this probably compares with our emergency room finding, 13% in children following a bike injury, and should not have limited the impact of the intervention.[24] Finally, Ottawa is a relatively affluent community with 33% of households reporting an income of more than $50 000 in 1985. Price was identified in our follow-up as a minor deterrent compared with lack of interest and peer pressure.[25]

In a similar study conducted in an emergency room following bike injuries, our helmet promotion efforts were also ineffective.[26] Our experience contains a message of caution, as well as a challenge. Helmet promotion is now encouraged by medical societies. While such interventions may have an impact over time, we suggest the medical community seek additional strategies.[27] There is evidence from Australia, Canada, and the United States that community programs involving the media, schools, and youth groups are effective.[9,28,29] Physicians can play an active role in teaching parents and children that bikes are not just another toy and that helmets can go a long way in reducing the risks of head injury. However, helmets for children are still fairly innovative, and the promotional challenges suggest physicians would be more effective by participating in the design and implementation of multidisciplinary community campaigns. ☐

TABLE 2—Subject Baseline Characteristics

Characteristic	Intervention (n = 167)	Control (n = 172)	Total (N = 339)
No. of children	288	288	576
Age			
Mean (years)	9.62	9.30	9.44
SD	3.43	3.14	3.29
Range	5–17	5–17	5–17
Families owning one or more helmets for cycling (%)	24.6	27.9	26.3

Acknowledgments

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The promotional material used in this study is as follows:

1. "Be bike smart." The Easter Seal Society, 24 Ferrand Drive Don Mills, Ontario, Canada M3C 3N2.
2. "Sprocketman talks road sense." State of North Carolina Department of Transportation, PO Box 25201, Raleigh, NC 27611.

3. "A consumer's guide to bicycle helmets." Washington Area Bicyclist Association, 1015 31st Street NW, 3rd floor, Washington, DC 20007.

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Preventing and Postponing Death: Trends in Tennessee Infant Mortality

Joyce M. Piper, DrPH

ABSTRACT

Infants born in Tennessee between 1984 and 1988 experienced reduced mortality compared with those born in the 5 years previous. For all birthweight singleton infants, the reduction from 10.2 to 8.4 deaths per 1000 live births represents an 18% decline in mortality. The most impressive gain made was for very low birthweight (<1500 g) infants in the neonatal period, where mortality was reduced 24%. However, for these infants there was nearly a sixfold increase in the postneonatal mortality associated with prematurity-related causes. ***(Am J Public Health.*** **1991;81:1046–1048)**

Introduction

Evidence from the National Center for Health Statistics indicates that the rate of decline in infant mortality slowed during the 1980s.[1] Much of the progress made in the previous decade appears to be due to improvements in birthweight-specific mortality associated with care available at tertiary perinatal care facilities.[2–9]

One consequence of the impressive gains made by these facilities in preventing the deaths of very low birthweight (<1500 g) infants has been the postponement of some deaths from the neonatal to the postneonatal period.[10–13] These postponed deaths are thought to have contributed to the lack of decline in postneonatal mortality during the 1970s. Little is known about this effect in more recent data.

To determine whether reductions in Tennessee infant mortality have been achieved during the 1980s and to quantify any evidence of a postponement effect, we reviewed vital records data for 1979 through 1988.

Joyce M. Piper is with the Department of Preventive Medicine, Vanderbilt University School of Medicine, Nashville, Tenn.

Requests for reprints should be sent to Joyce M. Piper, DrPH, Department of Preventive Medicine, Nashville, TN 37232-2637.

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TABLE 1—Mortality Rates per 1000 Live Births By Birthweight Category for Tennessee Infants, 1979–1988

Birthweight Category	Mortality Rate: Infants Born 1979–1983	Mortality Rate: Infants Born 1984–1988	Difference in Rates (later vs earlier)	95% Confidence Interval
NBW(≥2500 g)				
Infant	5.0	4.0	−1.0	−1.3 to −0.7
Neonatal	2.1	1.6	−0.5	−1.8 to −0.3
Postneonatal	2.9	2.4	−0.5	−0.8 to −0.2
n	304 765	302 194	—	—
MLBW(1500–2499 g)				
Infant	31.1	26.9	−4.2	−7.6 to −0.8
Neonatal	19.8	16.4	−3.4	−6.1 to −0.7
Postneonatal	11.3	10.5	−0.8	−2.9 to 1.3
n	19 155	18 672	—	—
VLBW(<1500 g)				
Infant	372.7	296.8	−75.8	−98.4 to −53.4
Neonatal	332.4	251.9	−80.5	−102.2 to −58.8
Postneonatal	40.3	44.8	4.5	−5.1 to 14.1
n	3249	3481	—	—
Total				
Infant	10.2	8.4	−1.8	−2.3 to −1.3
Neonatal	6.4	5.1	−1.3	−1.7 to −0.9
Postneonatal	3.8	3.3	−0.5	−0.8 to −0.2
n	327 169	324 347	—	—

Note: NBW = normal birthweight; MLBW = moderately low birthweight; VLBW = very low birthweight.

Materials and Methods

Design, Population, and Data Sources

Two birth cohorts were studied: infants born in the period 1979 to 1983 and 1984 through 1988. Infants were grouped into two 5-year periods to dampen year-to-year rate fluctuations. Criteria for study entry included singleton live birth during 1979 through 1988, a birthweight of 500 to 5999 g, maternal race either White or Black, and Tennessee residence at birth and, for those infants who died, a death with recording of the vital events in Tennessee.

Infant death certificates were matched to corresponding birth certificates using a computerized linkage algorithm. Overall, birth information was available for 97.0% of the infants who died from the 1979 to 1983 birth cohort and for 96.7% of the infants who died from the later birth cohort. Investigation of a sample of the unlinked infant death certificates revealed that most were attributable to infant surname changes.

Study Definitions

Infants were classified into three birthweight categories: normal birthweight (NBW = ≥2500 g), moderately low birthweight (MLBW = 1500 to 2499 g), and very low birth weight (VLBW = <1500 g).

Analysis

The study analysis compared mortality rates for the two birth cohorts. Rate differences were calculated by subtracting rates of the 1979 to 1983 cohort from those of the 1984 to 1988 cohort. Confidence intervals (CI) of differences were calculated with standard techniques.[14]

We aggregated infant deaths into eight leading cause of death categories, as suggested by Dollfus et al.[15] This categorization scheme uses underlying cause, as reported on death certificates and coded with the International Classification of Disease, Ninth Revision (ICD-9), to group infant deaths that have similar etiologies or strategies for prevention.[15]

Results

In each of the three birthweight categories (NBW, MLBW, VLBW) there were reductions in the infant mortality rate for the later birth cohort, and these reductions were most pronounced for VLBW infants (Table 1). There were also reductions in neonatal mortality rates for all three groups, the greatest reduction occurring for the VLBW infants. However, reduction in postneonatal mortality was limited to NBW infants.

To examine these results further, we classified infant deaths for VLBW and MLBW infants using the categorization scheme previously described. For VLBW infants, there were reductions in the neonatal mortality rate for deaths associated with prematurity, birth asphyxia, and obstetric conditions (Table 2). However, for these infants, the postneonatal mortality rate for deaths associated with prematurity increased from 3.1 to 17.8 per 1000 live births. For MLBW infants, there were reductions in the neonatal mortality rate associated with prematurity (rate difference = −1.8, 95% CI of the difference −2.3–−1.3) and perinatal infections (rate difference = −2.1, 95% CI of the difference −3.3–−0.9). For these infants, there were no changes in the causes of postneonatal mortality.

Discussion

Our data show reduced infant mortality for infants born between 1984 and 1988 compared with those born in the five years previous. For infants of all birthweights the reduction from 10.2 to 8.4 deaths per 1000 live births represents an 18% decline in infant mortality. The most impressive gain made was for VLBW infants in the neonatal period, where the mortality rate declined from 332.4 to 251.9 deaths per 1000 live births, a 24% decrease. There was a small improvement in postneonatal mortality overall, but no improvement in this parameter for either MLBW or VLBW infants. For VLBW infants, there was nearly a sixfold increase in postneonatal mortality associated with prematurity-related deaths, suggesting that more of these fragile newborns are surviving the neonatal period only to die later of the complications of preterm birth.

It is unlikely that ascertainment bias could account for our findings, because we were able to link the same proportion of infant deaths to their birth certificates in the two time periods. The computerized scheme we used to group infant deaths of related causes was not subject to differential investigator surveillance. However, it is possible that changes in the attribution of underlying cause of death over time could have resulted in the increase of prematurity-related postneonatal deaths observed.

The postponement of the deaths of some VLBW infants from the neonatal to the postneonatal time period remains evident when data from the middle to late 1980s are compared with data from the early 1980s. These postponed deaths

TABLE 2—Neonatal and Postneonatal Mortality Rates by Cause of Death for Very Low Birthweight (VLBW) Infants Born in Tennessee, 1979 to 1988

Cause of Death	VLBW Infants Born in 1979–1983		VLBW Infants Born in 1984–1988		Difference in Rates (later vs earlier)	95% Confidence Interval
	Rate	(n)	Rate	(n)		
Neonatal						
Prematurity related	241.9	(786)	178.4	(621)	−63.5	−83.0 to −44.0
Perinatal infections	15.4	(50)	13.2	(46)	−2.2	−7.9 to 3.5
Birth asphyxia	15.7	(51)	9.2	(32)	−6.5	−11.9 to −1.2
Obstetric conditions	13.9	(45)	7.8	(27)	−6.1	−11.1 to −1.1
Congenital anomalies	20.9	(68)	19.2	(67)	−1.7	−8.4 to 5.0
SIDS[a]	—	(0)	—	(0)	—	—
Infections (nonperinatal)	2.8	(9)	2.9	(10)	0.1	−2.4 to 2.6
Injuries	0.3	(1)	0.3	(1)	0	−0.8 to 0.8
Other	21.5	(70)	21.0	(73)	−0.5	−7.4 to 6.4
Postneonatal						
Prematurity related	3.1	(10)	17.8	(62)	14.7	9.9 to 19.5
Perinatal infections	—	(0)	0.6	(2)	0.6	−0.2 to 1.4
Birth asphyxia	0.3	(1)	—	(0)	−0.3	−0.8 to 0.3
Obstetric conditions	—	(0)	—	(0)	—	—
Congenital anomalies	3.1	(10)	3.4	(12)	0.3	−2.4 to 3.0
SIDS[a]	6.2	(20)	4.9	(17)	−1.3	−4.8 to 2.2
Infections (nonperinatal)	9.2	(30)	5.7	(20)	−3.5	−7.6 to 0.6
Injuries	0.9	(3)	—	(0)	−0.9	−1.9 to 0.1
Other	17.5	(57)	12.3	(43)	−5.2	−17.6 to 7.2

[a] Sudden infant death syndrome.

probably represent an inevitable consequence of the application of technologically advanced medicine to multiple problems of high-risk newborns. Those concerned with improving infant health need to continue to examine both neonatal and postneonatal mortality, because early survival alone may not provide a complete picture. □

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A Comparison of Breast-Feeding Data from the National Surveys of Family Growth and the Ross Laboratories Mothers Surveys

Alan S. Ryan, PhD, William F. Pratt, PhD, Jeffrey L. Wysong, MS, Gregory Lewandowski, MS, James W. McNally, MA, and Fritz W. Krieger, MS

ABSTRACT

This study compares rates of breast feeding reported in the National Surveys of Family Growth with those from the Ross Laboratories Mothers Surveys. Both surveys have documented rates of breast feeding over the last 30 or more years. Despite differences in survey methodology, both surveys document similar long-term trends in breast feeding. The similarities of rates in breast feeding also are evident across several maternal sociodemographic characteristics. We conclude that both surveys produce reliable and useful estimates of breast feeding. (*Am J Public Health.* 1991; 81:1049–1052)

Introduction

The principal sources of information on breast-feeding trends in the United States are the National Surveys of Family Growth (NSFG) and the Ross Laboratories Mothers Surveys (RLMS). These two surveys have documented trends in breast feeding over the last three decades. This study compares reported rates of breast feeding in the NSFG with those in the RLMS between 1955 and 1987 and compares the rates in 1987 according to selected sociodemographic characteristics.

We found that despite differences in survey methodologies and design, both surveys document similar trends in breast-feeding, and these similarities are evident across sociodemographic characteristics. The similarities of breast-feeding rates bear witness to the usefulness of the surveys in documenting the incidences and long-term trends of breast-feeding practices in the United States.

Methods

National Surveys of Family Growth

Information on breast-feeding has been collected by the National Center for Health Statistics in four NSFGs: Cycle I (1973), Cycle II (1976), Cycle III (1982), and Cycle IV (1988).[1–5] The surveys were designed to provide information on factors affecting childbearing, such as contraception, sterilization, infertility, and related aspects of maternal and infant health.[6] In addition to a series of detailed demographic questions, during a face-to-face personal interview, each mother was asked, "When your child was an infant, did you breast-feed (him/her) at all?" (i.e., ever-breast-fed).

Because the surveys focused on fertility, the NSFG included samples that were representative of the national population of civilian noninstitutionalized women in their reproductive years (ages 15 through 44 years). Cycles I and II had sample sizes of 9800 and 8600 women, respectively, who were currently or previously married. Cycle III included 7969 women of all marital statuses and women living in group quarters. The Cycle IV target population was expanded to include women living in Alaska and Hawaii. In Cycle IV, 8450 women were selected; this sample was drawn from households interviewed between October 1985 and March 1987 for another NCHS survey, the National Health Interview Survey (NHIS). The response rate to the NHIS was 96%, and the response rate to Cycle IV of the NSFG was 82% of the women subsampled from the NHIS. Thus, the total NSFG Cycle IV response rate is approximately 79%.

Ross Laboratories Mothers Surveys

Ongoing surveys by Ross Laboratories have documented trends in breast feeding in the United States since 1955.[7–12] The RLMS is a large, national mail survey designed to determine patterns of infant feeding during the first 6 months of life.

Each quarter, questionnaires are mailed to mothers when their infants reach 6 months of age. Mothers are asked to recall the type of milk fed to their infant immediately after birth, typically in the hospital, and during each month of the first 6 months of life.

The sample is derived from a list of births that consists of approximately 70% to 82% of all new mothers in the United States. The list includes names from hospital sources, county records of birth reg-

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Alan S. Ryan, Jeffrey L. Wysong, Gregory Lewandowski, and Fritz W. Krieger are with Ross Laboratories, Columbus, Ohio. Alan S. Ryan is also associated with the Department of Anthropology of Ohio State University. William F. Pratt is with the National Center for Health Statistics, Hyattsville, Md. James W. McNally is a student at Brown University.

Requests for reprints should be sent to Alan S. Ryan, PhD, Ross Laboratories, 625 Cleveland Avenue, Columbus, OH 43216.

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TABLE 1—Incidence of Breast Feeding, 1955–1987. National Surveys of Family Growth (NSFG) Compared with Ross Laboratories Mothers Survey (RLMS)

	NSFG	RLMS	
Year	Breast Fed at Any Time	In Hospital	1 Week
1955–1959	32.0	—	28.7
1960–1964	27.0	—	27.4
1965–1969	25.0	28.1	26.0
1970	23.0	26.5	24.9
1971	25.0	24.7	22.8
1972	22.0	28.1	25.9
1973	25.0	28.9	26.8
1974	30.0	32.2	29.6
1975	34.0	35.5	33.4
1976–1977	42.2	43.2	40.9
1978–1979	44.3	48.8	47.4
1980–1981	52.5	56.4	55.2
1981	57.9	57.6	56.4
1982	59.1	61.9	60.5
1983	55.3	58.4	57.0
1984[a]	54.2	59.2	57.6
1985[a]	55.6	57.3	55.6
1986[a]	53.5	56.3	54.9
1987[a]	56.3	56.1	54.7

[a] RLMS breast-feeding data have been adjusted to reflect time that survey was mailed (adjusted year = last half of year + first half of following year; see text).

istrations, photography and entrepreneurial services, and newspapers.

Throughout the past few years, the number of questionnaires mailed to new mothers has increased substantially. In 1984, 56 894 questionnaires were mailed and 30 694 (54%), or 0.8% of annual births, were returned. In 1987, 154 958 questionnaires were mailed and 83 240 (54%), or 4.1% of annual births, were returned.

The collected infant-feeding data are then weighted by subclass means in the sample to account for differing responses and coverage rates. The subclasses were defined by the variables that have historically been associated with breast-feeding: maternal region of residence, ethnic background, age, and education. The weights were assigned using continually updated information from the US Bureau of the Census and National Center for Health Statistics.[13–16]

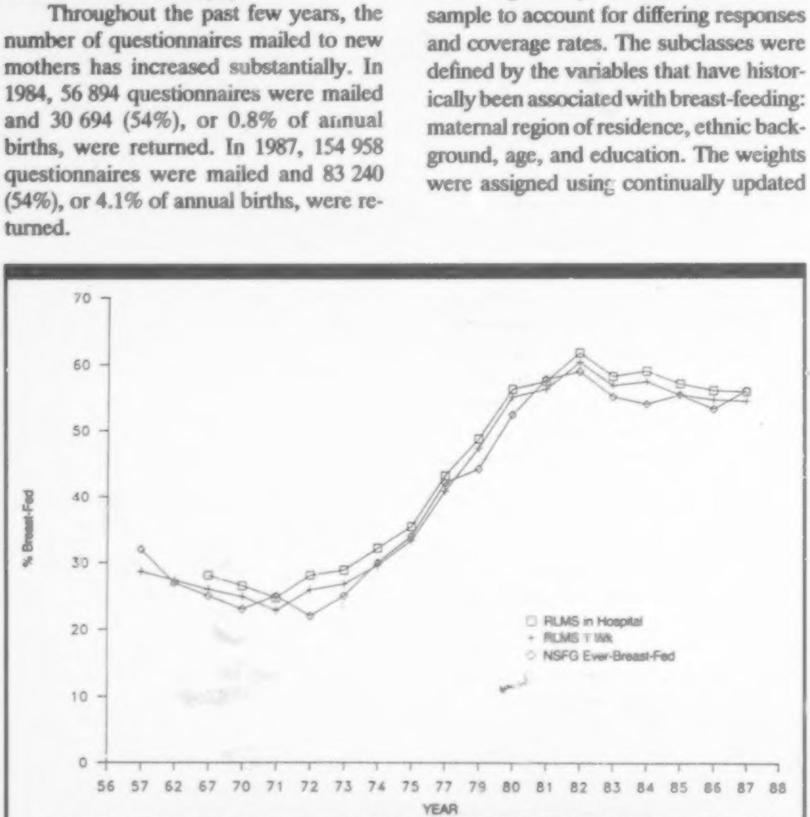

FIGURE 1—Incidence of breast feeding, 1957–1987: National Surveys of Family Growth (NSFG) compared with Ross Laboratories Mothers Survey (RLMS). The term "breast feeding" is defined as the exclusive use of human milk or the use of human milk with a supplemental bottle (i.e., formula or cow's milk).

Data Analyses

Breast-feeding rates reported at two time periods in the RLMS (in the hospital and at 1 week of age) were compared with the ever-breast-fed rates reported in the NSFG. We also considered rates of breast feeding in 1987 (most recent data in both surveys) according to the following sociodemographic characteristics: maternal age, parity, ethnic background, education, region of residence, and birthweight.

Because the RLMS questionnaire is mailed to mothers when their infants are 6 months of age, to produce yearly estimates that are comparable to those provided in the NSFG, it was necessary to adjust for the time the RLMS survey was mailed. For example, the adjusted data for 1987 represent information collected during the last half of 1987 and the first half of 1988 (Table 1). These adjusted data may differ slightly from previously published data that have not been adjusted.[11]

Since the data in the RLMS and NSFG were obtained through complex sample designs, it was necessary to use specific statistical techniques that incorporated the sample weights for the computation of estimated breast-feeding percentages and standard errors.[6,17,18] Differences between rates of breast-feeding in the two surveys were tested for statistical significance using a z test ($P<.05$).

Results: Trends in Breast Feeding

Both surveys show a small decline in the incidence of breast feeding from 1955 to the early 1970s, followed by an increase until 1982, when the incidence of breast feeding reached its peak (Figure 1). Between 1957 and 1988, the two surveys demonstrate an approximate twofold net increase in the incidence of breast feeding (32.0% to 56.3% in NSFG; 28.7% to 54.7% in RLMS at 1 week of age). Since 1982, as reported in RLMS, the incidence in the initiation of breast feeding and breast feeding at 1 week of age steadily declined (61.9% to 56.1% in the hospital; 60.5% to 54.7% at 1 week of age). The NSFG ever-breast-fed data showed a statistically nonsignificant net decline between 1982 and 1987 (59.1% to 56.3%).

Since 1955, at each year, differences between the two surveys in rates of breast-feeding were not large. Over time, the av-

TABLE 2—Incidence of Breast-Feeding: Ross Laboratories Mothers Survey (RLMS) Compared with 1988 National Survey of Family Growth (NSFG) for 1987

	RLMS[a]		NSFG[b]				
Characteristics	N	Breast Feeding (%)	N	Breast Feeding (%)	Difference	Standard Error of Difference	95% Confidence Interval of Difference
Total United States	83 240	54.7	563	56.3	−1.6	3.2	−7.9, 4.7
Ethnicity[d]							
Black (non-Hispanic)	10 151	25.3	195	23.1	2.2	4.0	−5.6, 10.0
White (non-Hispanic)	66 136	61.1	338	61.0	0.1	3.5	−6.8, 7.0
Hispanic	5 423	49.7	50	41.8	7.9	8.1	−8.0, 23.8
Maternal education[d]							
High school or less	36 043	43.9	334	44.6	−0.7	4.6	−9.7, 8.3
Some college	46 816	72.0	229	71.6	0.4	3.5	−6.5, 7.3
Maternal age (years)[d]							
<20	4 544	30.5	59	26.1	4.4	6.6	−8.5, 17.3
20–24	19 389	47.1	168	50.2	−3.1	7.3	−17.4, 11.2
25–29	31 425	60.9	185	60.1	0.8	4.1	−7.2, 8.8
30–34	20 677	66.4	104	61.5	4.9	5.6	−6.1, 15.9
>34	7 088	66.7	47	78.1	−11.4	6.9	−24.9, 2.1
Region							
Northeast	14 976	51.6	110	68.3	−16.7	5.9	−28.3, −5.1[c]
South	27 761	46.0	201	41.1	4.9	5.2	−5.3, 15.1
Midwest	21 720	52.6	137	48.6	4.0	4.4	−4.6, 12.6
West	18 783	72.9	115	74.1	−1.2	5.3	−11.6, 9.2
Residence							
Urban	62 975	55.3	449	57.4	−2.1	3.7	−9.4, 5.2
Rural	20 265	52.9	114	52.9	0.0	5.2	−10.2, 10.2
Parity[d]							
Primiparous	41 098	55.3	188	56.3	−1.0	4.7	−10.2, 8.2
Multiparous	41 028	54.2	375	56.3	−2.1	4.3	−10.5, 6.3
Birthweight (g)[d]							
<2500	4 113	38.4	43	32.7	5.7	10.1	−14.1, 25.5
>2500	75 649	56.1	518	57.7	−1.6	3.3	−8.1, 4.9

[a] Breast feeding reported at 1 week of age.

[b] Infants who have been breast fed at any time.

[c] $P<.05$.

[d] The sample sizes for these demographic subsets do not add to total because of unit nonresponse.

erage difference between the NSFG ever-breast-fed and the RLMS in-hospital breast-feeding rates was 2.6%; differences ranged from 0.3% in 1971 to −6.1% in 1972. At each year, the RLMS in-hospital breast-feeding rates tended to be higher than the NSFG ever-breast-fed rates.

Differences between the NSFG ever-breast-fed and the RLMS breast feeding at 1 week of age rates were smaller than differences between the NSFG ever-breast-fed and RLMS in-hospital rates; the average difference was only 0.6%, and differences ranged from 3.3% in the period 1955 to 1959 to −3.9% in 1972 (Figure 1).

As shown in Table 1, both surveys demonstrated that in 1987, breast-feeding was relatively high among women who were White, older in age, college educated, living in the urban and western regions of the United States, and who had an infant of normal birthweight. Also, according to both surveys, the lowest rates of breast feeding were among women who were Black, with less than a college education, residing in the rural and southern regions of the country, younger in age, and who had an infant of low birthweight.

The *z*-test analyses indicated that across the socioeconomic subgroups considered, except for the northeast region, differences in breast-feeding rates between the two surveys were not statistically significant (Table 2).

Discussion

The present study indicates that over the last 30 or more years, differences in breast-feeding rates between the NSFG and RLMS surveys were small. Both surveys agree in showing a decline in breast-feeding from 1955 to the early 1970s, followed by an upward trend until 1982. They also agree in demonstrating that, across several selected maternal characteristics, certain subgroups of mothers are more likely to breast-feed than others.

The NSFG ever-breast-fed rates and the RLMS breast feeding at 1 week of age rates were comparable (Table 1). The RLMS in-hospital breast-feeding rates averaged approximately 2% to 3% higher than the NSFG ever-breast-fed rates. In the NSFG, mothers were asked if they ever breast-fed their infant, whereas in the RLMS, mothers were asked to indicate the kinds of milk they fed to their infant. A possible explanation for the lower in-hospital breast-feeding rates in the NSFG is that respondents who fed breast milk to their infant in the hospital but not after leaving may have considered such a brief encounter a nonexperience. In contrast, mothers in RLMS were not asked if they actually "breast-fed"; they may have given their infants a bottle containing expressed human milk. Because the questions used to define breast-feeding were not strictly comparable, the observed differences in the in-hospital breast-feeding rates may reflect variation in the manner in which mothers define breast-feeding.

Except for breast-feeding in the northeast region of the United States, across all other sociodemographic subgroups of mothers, differences in breast-feeding rates between the two surveys were not statistically significant. Considering the many comparisons that were made between these two surveys on mothers' characteristics, the finding of a single, statistically significant difference is not surprising and probably warrants little attention.

The principal sources of data on long-term trends in breast-feeding are the National Surveys of Family Growth and the Ross Laboratories Mothers Survey. The similarities in estimates support the usefulness of these surveys for documenting rates of breast-feeding in the United States. ☐

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Risk Factors for Childhood Homicides in Ohio: A Birth Certificate–Based Case-Control Study

Kim A. Winpisinger, MS, Richard S. Hopkins, MD, MSPH,
Robert W. Indian, MS, and Jeptha R. Hostetler, PhD

ABSTRACT

Death certificates of children less than 8 years of age who were killed between 1979 and 1986 were linked to their Ohio birth certificates and compared with those of Ohio children born in 1983 (controls). Having an unmarried mother increased risk of homicide almost fivefold (odds ratio 4.87). Having a teenage mother, a mother who had not graduated from high-school, and being of Black race or low birthweight each increased the risk by approximately threefold. Increases in the proportion of children born to unmarried mothers may contribute to increases in childhood homicide rates. (*Am J Public Health.* 1991;81:1052–1054)

Introduction

Rates of death by homicide in children have increased markedly in the United States and in Ohio in the last 30 years, particularly for children under age 5.[1–3] While many of the risks associated with childhood homicide are known,[4–11] we undertook a case-control study of homicides in Ohio children under the age of 8, to identify risk factors and to clarify the associations among those factors.

Methods

Cases were taken from death certificates of child homicide cases (International Classification of Disease, revision 9 [ICD-9] codes E-960 through E-969) from the years 1979 through 1986, of Ohio residents who were born during those same years. For each case, the birth certificate was sought in the Ohio birth certificate file, matching on last name, first three letters of first name, and date of birth. Con-

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Kim A Winpisinger, Richard S. Hopkins, and Robert W. Indian are with the Ohio Department of Health, Columbus, Ohio. Jeptha R. Hostetler and Richard S. Hopkins are with the Ohio State University College of Medicine, Columbus, Ohio.

Requests for reprints should be sent to Ms. Winpisinger, Bureau of Epidemiology and Toxicology, Ohio Department of Health, 246 N High Street, PO Box 118, Columbus, OH 43266-0118.

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TABLE 1—Characteristics of Ohio Child Homicide Cases, Ages 0 to 7 Years, 1979–1986, and 1983 Ohio Birth Cohort Controls

Characteristic	Case	Control
Born out of wedlock (%)	54.2	19.6
Black race (%)	38.0	15.1
Low birthweight (%)	16.2	6.7
Mother not a high school graduate (%)	47.4	20.8
Pregnancy complications (%)	12.0	9.6
Illness in pregnancy (%)	4.2	3.4
Singleton birth (%)	96.5	97.9
Male (%)	60.6	51.2
Mother's age (mean years)	21.9	25.3

TABLE 2—Odds Ratios (ORs) and Confidence Intervals (CI) for Selected Risk Factors for Child Homicide in Ohio Children, Ages 0 to 7 Years, 1979–1986, Compared with 1983 Ohio Births as Controls

Risk Factor	OR[a]	95% CI[b]
Born out of wedlock	4.87[a]	3.45–6.89
Black race	3.52[a]	2.47–5.02
Low birthweight	2.75[a]	1.71–4.38
Mother not a high-school graduate	3.45[a]	2.43–4.88
Pregnancy complications	1.32	0.77–2.24
Illness in pregnancy	1.29	0.51–3.02[b]
Singleton birth	0.57	0.22–1.57[b]
Labor complications	1.15	0.78–1.68
Metro county resident	1.52[a]	1.04–2.21
Congenital anomalies	0.55	0.03–3.64[b]
Teenaged mother	3.20[a]	2.23–4.60

[a] Statistically significant

[b] Less than five frequencies per cell

trols, also obtained from Ohio birth certificates, were all children born live to Ohio residents in 1983.

Factors examined in this study included marital status of the victim's mother at the time of the victim's birth, race (Black vs White), educational attainment of the mother (<12 years vs ≥12 years), county type (metropolitan vs nonmetropolitan), birthweight (<2500 g vs ≥2500 g), mother's age (teenager or ≥20), presence or absence of birth complications, illness during pregnancy, labor complications, multiple birth, and congenital anomalies.

Odds ratios (ORs) were calculated for each of these potential risk factors. Those risk factors associated with the outcome in a univariate analysis were examined in a stratified analysis, two at a time. Cornfield's method was used to calculate 95% confidence limits around the odds ratios.[12]

Results

There were 189 children who were born and died during the study period; birth certificates were identified for 140, or almost 75%. Of these, 134, or 96%, were under age 5. Mean age at death was 1.1 years. Children for whom birth certificates could not be found were more likely to be Black (48% vs 39% for whom certificates were found), metropolitan residents (78% vs 30%), and/or less than 28 days old at the time of death (29% vs 7%).

Cases were more likely to be Black, male, and of low birthweight than controls (Table 1). Likewise, case mothers were younger at the time of the child's birth and less likely to be married or to have a high school diploma.

Univariate ORs (Table 2) were highest for being born out of wedlock, Black race, having a mother who did not finish

TABLE 3—Crude and Stratified Odds Ratios (ORs) for Homicide, Ohio Cases under 7 Years of Age from 1979 to 1986, Compared with All 1983 Ohio Births

Risk Factor	Out-of-Wedlock Birth	Black	Teenaged Mother	Metro County Residence	High School Nongraduate	Low Birthweight
Crude	4.87[a]	3.52[a]	3.20[a]	1.52[a]	3.45[a]	2.75[a]
Stratified by:						
Legitimacy						
Legitimate	—	1.51	3.24[a]	0.97	2.41[a]	2.77[a]
Out of wedlock	—	1.84[a]	1.27	1.47	1.82	1.77
Race						
Black	4.29[a]	—	1.45	2.44	2.86[a]	2.72[a]
White	3.84[a]	—	4.06[a]	0.98	3.14[a]	1.70
Mother's age						
<20 years	2.04[a]	1.48	—	1.28	1.68	0.85
≥20 years	5.19[a]	4.13[a]	—	1.65[a]	3.16[a]	3.95[a]
County type						
Metro	5.37[a]	3.54[a]	2.96[a]	—	4.65[a]	2.77[a]
Nonmetro	3.54[a]	1.42[a]	3.82[a]	—	1.78	2.16
Mother's education						
Not a high school graduate	4.41[a]	2.73[a]	1.58	2.68[a]	—	1.88
High school graduate	3.33[a]	2.98[a]	2.98[a]	1.03	—	2.70
Birthweight						
Low	3.10[a]	4.72[a]	0.84	1.72	2.49	—
Normal	4.86[a]	2.96[a]	3.89[a]	1.34	3.58[a]	—

[a] Lower bound of 95% confidence limit does not include 1.0.

TABLE 4—Homicide Deaths and Out-of-Wedlock Births of Ohio Children by Race and Specific Years

	No. of Homicide Deaths under Age 5 Years	Homicide Rate per 100 000 Deaths under Age 5 Years	All Ohioans born to unwed Mothers (%)
White			
1970	17	2.1	5.8
1980	19	2.8	11.0
1986	12	1.7	15.7
Black			
1970	7	7.2	38.3
1980	9	9.3	56.8
1986	20	19.9	64.0

high school, having a teenaged mother, having a birthweight less than 2500 g, and living in a metropolitan county.

The effect of having an unmarried mother on risk of homicide was strong in all subgroups examined (Table 3), but was greatest among children whose mothers were age 20 or older, and/or lived in a metropolitan area, and among children who were of normal birthweight.

The crude risk of homicide associated with Black race was 3.52 (Table 3), but the risk fell to between 1.5 and 1.8 when separate strata of children born to married and unmarried mothers were examined. Thus, the apparent effect of Black race on risk of homicide is largely explained by the fact that a larger proportion of Black infants were born to unmarried mothers. The effect of Black race on risk of homicide is greatest in children residing in metro counties, in low-birthweight children, and in children with mothers over age 20, and varies little by educational status of the mother.

The crude risk of dying from homicide is increased approximately 3.2-fold when the child's mother is a teenager at the time of the child's birth. The magnitude of this increase in risk is greatest when the child is otherwise at lower risk.

The crude OR for metropolitan residence is only of modest magnitude (1.52), but the increase in risk is much greater when the mother is not a high school graduate or when the mother is Black.

Children born to mothers who did not finish high school (a proxy for low socioeconomic status[13–15]) are approximately 3.45 times more likely to become homicide victims than are children born to high school graduates. This OR estimate appears to be confounded in part by effects of race, mother's age, and marital status, as the stratum-specific estimates for both strata of all three variables were noticeably less than the crude OR.

Low birthweight was a significant risk factor for homicide in most subgroups. Low birthweight is more strongly associated with risk of homicide among Black than White infants, and among children whose mothers were married and age 20 or older at the time of their child's birth.

Discussion

Many of the expected relationships found in other studies were observed in these data: children of younger, poorly educated mothers, Black children, children of low birthweight, children whose mothers were unmarried at the time of the child's birth, and children who resided in metropolitan areas were all more likely to be killed than other children. In general, these variables appear to be independently associated with the risk of homicide. Variables associated with the pregnancy were not good predictors for risk of homicide, nor were congenital malformations. The effect of race is largely explained by the high prevalence of Black infants with unmarried mothers. Similarly, the crude effect of low educational attainment of the mother on risk of homicide appears to be explained in part by the fact that these mothers are more likely to be unmarried. A weaker effect of having an unmarried or teenaged mother on risk of homicide was seen when the infant was otherwise at lower risk. This suggests that other social supports in the child's environment may reduce the risk of homicide.

The consistent association of homicide risk with being born to an unmarried mother is striking in these data. It is unclear whether this finding represents a causal relationship, or whether both increased risk of homicide and being born to an unmarried mother are consequences of some other risk factor that was not captured in our study. If there is a causal relationship, we can expect the incidence of childhood homicide to continue to increase, because the percentage of Ohio children born to unmarried mothers has increased dramatically in recent years (Table 4).

Children have many health and social problems, which makes the causes of childhood homicide complex. However, interventions to help teenagers finish high school, begin careers, and find permanent life partners before becoming parents would probably help prevent childhood homicide and abuse.[16] Such interventions are, in any case, desirable for other reasons. The strong associations between having a young or unmarried mother and the risk of homicide suggest that interventions to prevent homicides in children should focus on these social variables. ☐

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Costs and Benefits of Implementing Child Survival Services at a Private Mining Company in Peru

Karen G. Foreit, PhD, Delia Haustein, MD, Max Winterhalter, MD, and Ernesto La Mata, MD

ABSTRACT

Costs and savings of child health services were studied in a private mining company in Peru. Despite considerable outlays for medical services, few children under age 5 were vaccinated, and half of their illnesses went untreated. Children who were attended at the company clinic usually received unnecessary medication. As a result of the study, the company hired additional staff to provide integrated maternal-child preventive health care and family planning and contracted for intensive training and periodic on-site supervision. In less than 2 years, vaccination coverage reached 75%, and virtually all children under age 1 were enrolled in growth monitoring. Prescriptions were reduced by 24%, including a 67% drop in antimicrobials. The cost of the new services was $13 200 for the first 2 years. Approximately $6800 has been saved in pharmaceuticals prescribed for respiratory infection and diarrhea. Recently, two more mines adopted maternal and child health and family planning services. It is hoped that cost-benefit arguments will encourage other companies to incorporate aggressive child survival measures into their health plans. (*Am J Public Health.* 1991;81;1055–1057)

Introduction

As public resources stagnate, developing countries must find alternative sources of financing for primary health. One underutilized resource is the commercial sector. Approximately 10% of the 1984 health expenditures in Peru were covered by payments made by employers,[1] a source that could be tapped for preventive health care.

Most commercial health plans only cover curative care. The present study examined whether expenditures for curative child health care could be reduced through preventive interventions and better prescription practices, while maintaining or improving quality of care. The study site was a privately owned Peruvian mine that provides medical coverage to approximately 800 workers and 4000 dependents, including 900 children under age 5.

Despite company investments of $1.7 million in welfare and medical care, the population showed high morbidity. A prospective cost-benefit analysis found that family planning and child survival services could improve health status, at a cost that would be compensated by savings in other services averted. The company implemented the recommended services; this report summarizes the findings of the child survival component.

Methods

Baseline conditions were assessed during 1986 and 1987. Data included company records and household and clinic surveys. The household survey was conducted in July 1986, with a random sample of 220 miners' wives. The clinic survey was conducted in March 1987 and included 132 consultations with children under age 5. Preventive health services were added in 1988. A follow-up conducted in late 1989 included service statistics, a campwide census, and a clinical survey that captured 206 visits by children under age 5. Wholesale costs of medications prescribed were calculated at baseline and at follow-up.

Results

Baseline Conditions

Infant and child health were precarious at the start of this study. Infant mortality was estimated at 120 deaths per 1000 live births. Contributing factors included lack of well-baby care, low vaccination coverage, high incidence of diarrhea and respiratory infection, and overmedication and use of inappropriate medications.

Lack of well-baby care. All consultations surveyed were motivated by illness. Physicians reported that they had no time for well-baby care and that mothers would not use it if it were offered.

Vaccination. The household survey found that fewer than 5% of the children ages 1 to 4 were completely vaccinated. No child was vaccinated during an observed clinic visit. A quarter of the children over age 1 who had not been vaccinated spontaneously reported having had measles.

Morbidity. Of the children surveyed, 44% reported diarrhea and/or cough in the last 3 days; 24% reported diarrhea and 29% cough. Half the reported cases of diarrhea and/or cough did not seek treatment. Nevertheless, children under age 5, although they make up only 17% of the population, accounted for 50% of all clinic visits. Etiology was not determined, but the high incidence of blood and/or mucus in the stool accompanied by vomiting suggests shigellosis.[2] Respiratory infection accompanied by diarrhea is characteristic of rotovirus.

Overmedication. International norms suggest minimal medication for di-

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Karen G. Foreit is with The Futures Group, Washington, DC. Delia Haustein is with Asociacion Benefica Prisma, Lima, Peru. Max Winterhalter is with Instituto Marcelino, Lima, Peru. Ernesto La Mata is with Milpo Mining Company, Lima, Peru.

Requests for reprints should be sent to Karen G. Foreit, PhD, Senior Research Analyst, The Futures Group, 1101 Fourteenth Street, NW, Washington, DC 20005-5601.

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TABLE 1—Mean Number of Medications Prescribed by Physician's Diagnosis for Children under Age 5, 1987 vs 1989

	Diagnosis							
	ARI[a] Alone		Diarrhea Alone		ARI and Diarrhea		Others	
Medication	1987	1989	1987	1989	1987	1989	1987	1989
Antimicrobials	1.3	0.6	1.8	0.4	1.4	0.3	0.7	0.5
Symptomatics	1.8	2.0	1.2	1.4	2.1	2.0	1.0	0.8
Others	0.7	0.2	0.5	0.6	0.5	0.4	1.5	0.8
Total medications	3.7	2.9	3.5	2.4	4.0	2.7	3.2	2.0
No. of cases	47	85	24	44	28	14	21	21

[a]ARI = acute respiratory infection.

TABLE 2—Projected and Actual Costs and Estimated Cost Savings of Child Survival Activities

		Actual Costs		Estimated Savings	
Item	Projected Costs	Year 1	Year 2	Year 1	Year 2
Training and supervision	1500	1657	1946		
Salary and benefits	6300	2807	6666		
Equipment and materials	265	154	0		
Total	8065	4618	8612	1618	5147

arrhea and respiratory infection: oral rehydration and ampicillin for diarrhea with blood or mucus, acetaminophen and benzathine penicillin for moderate respiratory infection, etc. Nevertheless, every patient under age 5 received at least 2 and on average 3.7 medications, regardless of symptoms. Prescriptions for respiratory infection and diarrhea exceeded international recommendations (see Table 1). Prescriptions were classified according to the *Physician's Desk Reference*[3] and its Peruvian counterpart, *Diccionario de Especialidades Farmaceuticas*.[4]

Intervention

The study recommended that the mine hire an additional physician and nurse to provide well-baby and maternal care, and follow World Health Organization norms[5] for treating acute respiratory infection (ARI) and diarrhea. First-year costs were estimated at $8065; projected savings in pharmaceuticals ranged from $10 600 to $12 600 per year.

The company hired extra staff and contracted for training and on-site supervision. Services began in May 1988, in coordination with a national vaccination campaign. During the first quarter of 1989, standardized diagnostic and treatment protocols for ARI and diarrhea were adopted, and management contracted for a second year of supervision.

Program Impact

Vaccination coverage. The census found 75% coverage among children ages 1 to 4. Virtually all children under age 1 had been appropriately vaccinated.

Well-baby care. Of all clinic visits made in November 1989, 16% were for well-baby care, compared with none at baseline. Approximately 90% of children under age 1 and 35% of children ages 1 to 4 had been enrolled in growth monitoring. Thirty-nine children with second- and third-degree malnutrition were enrolled in nutritional rehabilitation; 31% returned to normal.

Treatment of diarrhea and respiratory infection. During 1989 the clinic treated 2258 cases of respiratory infecion and 1233 cases of diarrhea among children under age 5. As shown in Table 1, physicians prescribed on average one fewer medication per consultation; mean number of antimicrobials prescribed for ARI and diarrhea decreased from 1.5 to 05.

Costs and Savings

During the first 2 years, the company invested $13 200 in new child services at a cost per capita of less than $7.50 per year, compared with baseline per capita expenditures of $160 for curative care. After 18 months, annual prescription savings were approximately $5000 (January 1990 prices). Assuming linear trends, savings during the first 2 years totaled $6800 without considering improved health status due to better vaccination and nutrition practices. Table 2 presents program costs and estimated savings.

Discussion

Most child survival efforts focus on children outside the formal health system. However, many children inside the system also do not have access to preventive care. Furthermore, access to curative services does not guarantee use of services or that appropriate treatment will be given if sought. Consequently, even children covered by private plans may suffer from poor health, and company costs are inflated by improper prescription practices.

The present study demonstrates that preventive health can be provided at modest costs and achieve significant savings. Mothers accepted well-baby services and had no complaints about reduced medications. The advantages to the employer of preventive health care are clear. In 1989, two more mines initiated services. If the model is replicated throughout the mining community, more than 200 000 children in the most remote locations in Peru will benefit. ☐

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Measles Vaccination Coverage among Latino Children Aged 12 to 59 Months in Los Angeles County: A Household Survey

Donnell P. Ewert, MPH, J. Conley Thomas, PhD, MPH, Lorraine Y. Chun, MPH, Robert C. Enguidanos, MPH, Stephen H. Waterman, MD, MPH

ABSTRACT

A high rate of measles transmission has occurred in Los Angeles County since 1987. To determine measles vaccination coverage among Latino preschoolers, the group at highest risk of measles during this period, a cluster survey was conducted among Latino households with children aged 12 to 59 months. Of 279 children enrolled, 240 (86%, 95% confidence interval = 81% to 91%) reportedly had received the measles vaccine. Measles vaccination coverage among Latino preschoolers in Los Angeles County is insufficient to halt measles transmission within this subpopulation. (*Am J Public Health.* 1991;81:1057–1059)

Introduction

The incidence of measles in Los Angeles County (LAC) was less than 2.0 cases per 100 000 person-years from 1985 to 1987, but measles incidence increased to 7.3 in 1988[1] and 15.3 in 1989[2] (Table 1), with Latinos experiencing the highest rate among all ethnic groups during both years. Of the 1366 probable and confirmed measles cases occurring between January 1988 and June 1989, 601 (44%) were Latino preschoolers, the group at highest risk during this period.[2]

To promote early immunization and to reduce the number of children susceptible to measles, the LAC Department of Health Services lowered the recommended age of measles-mumps-rubella (MMR) vaccination from 15 to 12 months of age in July 1988.[3] During 1989, the same agency conducted a household cluster survey among Latino preschoolers to assess vaccination coverage.

Methods

The household survey was performed with a cluster sampling technique similar to that commonly used by the World Health Organization's Expanded Programme on Immunization (EPI).[4,5] The sampling frame included the 344 census tracts in LAC consisting of 50% or more Latino residents; two thirds of LAC Latinos reside within these census tracts. Census tracts were determined to be 50% or more Latino based on the 1980 census, adjusted for reported births and deaths through 1988.

Thirty census tracts were systematically selected from the sampling frame with probability proportionate to size. After random number tables and coin flips were used to select a street and initial household within each census tract, households were visited consecutively until seven eligible households were included in the survey. Households were not revisited if no one responded to a knock on the door.

Household inclusion criteria were permanent residence of a child aged 12 to 59 months who had at least one Latino parent, and presence of a respondent at least 16 years of age who was willing to participate. Latino was self-defined by the respondents and was generally understood to mean of Latin American descent. Possession of a written vaccination record

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This study was conducted at the Department of Health Services, Acute Communicable Disease Control, of the County of Los Angeles.

Requests for reprints should be sent to Donnell Ewert, MPH, Acute Communicable Disease Control, 313 N Figueroa Street, Room 231, Los Angeles, CA 90012.

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TABLE 1—Measles Incidence and Case Fatality Rates in Los Angeles County,[a] 1985 to 1989

	Measles			
Year	Cases	Incidence[b]	Deaths	Case Fatality Rate (%)
1985	44	0.60	0	0
1986	40	0.53	0	0
1987	126	1.66	0	0
1988	570	7.29	2	0.35
1989	1196	15.28	14	1.17

[a] Excluding Pasadena and Long Beach cities.
[b] Cases per 100 000 person-years.
Source: Communicable Disease Control Programs: Communicable disease morbidity reports, 1985–1989. County of Los Angeles Department of Health Services, Public Health Programs and Services.

was not required but was used if available. All eligible children in the household were included in the survey.

A standard interview form was used to obtain family information and vaccination history of the child(ren). A child was considered vaccinated against measles if the child had received either MMR or monovalent measles vaccine from any source on or after the first birthday. Interviewers were bilingual (English/Spanish) and conducted interviews between 9 AM and 5 PM on weekdays, from July 18 through September 13, 1989.

Age of vaccination was calculated from the birth and vaccination dates of children with written vaccination records. Confidence intervals (CIs) for the vaccination coverage point estimates were calculated using the method described by Lemeshow and Stroh.[6] Statistical significance was assessed with Mantel-Haensel χ^2 statistics, but these CIs may be too narrow because the survey did not employ a simple random sample.

Results

Of the 1684 households visited during the survey, an adult was present at 909 (54%), 676 of which were Latino. Of the 273 Latino households with eligible children, 209 (77%) agreed to participate. Fifty-one (80%) of the 64 refusals to participate were from secondary caretakers, who denied knowledge of vaccination history. The survey included 279 children, and the child's mother was the respondent for 221 (79%).

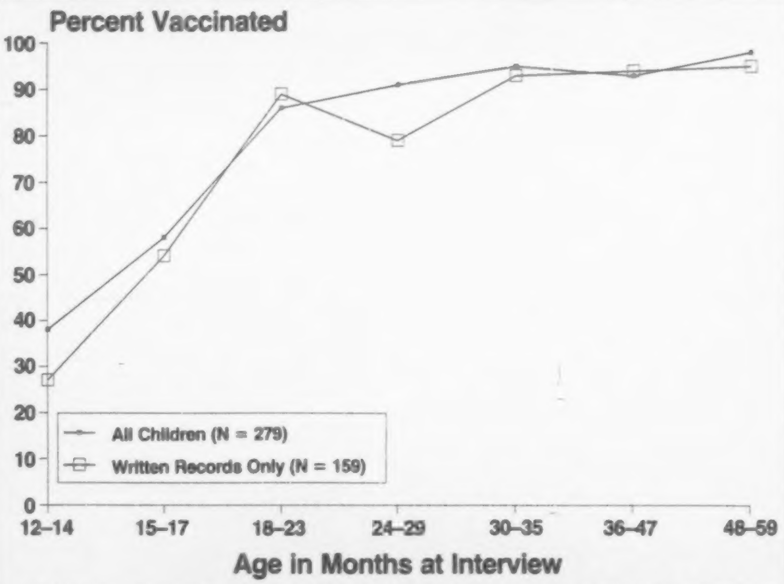

FIGURE 1—Proportion of Latino Children Residing in Predominantly Latino Census Tracts in Los Angeles County Who Were Vaccinated Against Measles, by Age Group, 1989

Respondents reported that 240 (86%; 95% CI = 81%–91%) of the children had received measles vaccine. Of the 201 children aged 24 to 59 months, 190 (95%) reportedly were vaccinated, compared with 50 (64%) of the 78 children aged 12 to 23 months (odds ratio [OR] = 9.7; 95% CI = 4.5–20.7) (Figure 1).

Written vaccine records were available for 159 (57%) of the children. Measles vaccine had been received by 131 (82%; 95% CI = 75% to 90%) of these children, a smaller proportion than that reported for children without written records (90%; 95% CI = 86–96%) (Figure 1). Mean age of measles vaccination was 19.5 months for the 131 vaccinated children, and mean age at interview was 21.5 months for the 28 unvaccinated children. Of the 46 children aged 15 to 29 months, 22 (48%) were vaccinated before 16 months of age, compared with 33 (34%) of the 98 children aged 30 to 59 months (OR = 1.8; 95% CI = 0.9–3.7) (Figure 2). Of the 112 children who were at least 2 years of age, 76 (68%) had been vaccinated before their second birthday.

Discussion

Effective measles vaccination programs aim to vaccinate a large enough proportion of susceptible persons (typically 92% to 96%) to attain herd immunity.[7] However, in nonrandomly mixed populations, such as the population of LAC, the number of susceptible persons and the frequency of their contact are more important in limiting transmission than the proportion of immune individuals within the whole population.[8] Because measles transmission was primarily focused in preschool children during 1989 and measles vaccination coverage was less than 92% to 96% among Latino preschoolers, endemic transmission will likely continue among Latinos in LAC until the number of susceptible preschool children decreases substantially, either through natural infection or vaccination.

Measles vaccination was delayed on the average until about 20 months of age among Latino children in this survey, and more than two thirds of the nonvaccinated children were younger than 24 months of age. These findings highlight the importance of delayed vaccination in the estab-

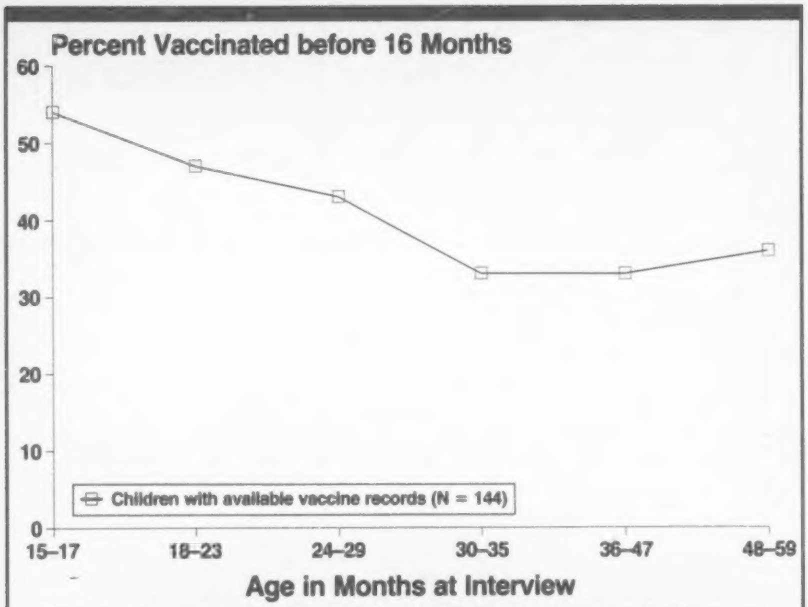

FIGURE 2—Proportion of Latino Children Residing in Predominantly Latino Census Tracts in Los Angeles County and for Whom Vaccine Records Were Available, Who Were Vaccinated Against Measles Before 16 Months of Age, by Age Group, 1989

lishment of a pool of children susceptible to measles. Fortunately, since the July 1988 change of recommended age of MMR vaccination from 15 to 12 months, the proportion of children vaccinated before 16 months of age has steadily increased (Figure 2).

The higher coverage reported among children without written records may reflect a tendency for the caretakers of these children to report vaccination for unvaccinated children. Inclusion of these children may have resulted in an overestimation of vaccination coverage. Also, children with mothers working outside the household may have been undersampled, since households were not revisited if no one answered the door or a knowledgeable respondent was not present. The effect of this potential bias on the coverage estimate is unknown.

Our sample variation from EPI—selection of seven households instead of seven children per cluster—had a negligible effect on the coverage estimate. An analysis that excluded the children who would not have been included in an EPI sample revealed a less than 1% difference in vaccination coverage. However, nonprobability sampling at the second stage, an inherent problem of the EPI-style survey, may have had some unknown effect on the vaccination coverage estimate.

Measles continues to affect primarily preschoolers in LAC, and the new two-dose MMR vaccination schedules recommended by the American Academy of Pediatrics[9] and the Immunization Practices Advisory Committee[10] will not reduce measles incidence among these children. Since about 95% of children receiving a second dose of MMR are already immune to measles,[11] public health efforts should focus on protection of unvaccinated preschoolers in areas where these children constitute the main risk group and resources are limited. ☐

Acknowledgments

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Low-Level Air Pollution and Upper Respiratory Infections in Children

Jouni J. K. Jaakkola, MD, Mikko Paunio, MD, Martti Virtanen, MD, and Olli P. Heinonen, MD, DSc

ABSTRACT

Effects of low-level air pollution were studied in one polluted city and two reference cities in northern Finland by comparing the frequency of upper respiratory infections over a 12-month period in 1982 as reported by parents of children ages 14 through 18 months (n = 679) and 6 years (n = 759). A similar comparison was carried out between children living in the more polluted and less polluted areas of the polluted city. The annual mean and the greatest half-hour concentrations of sulfur dioxide (23 and 807 $\mu g/m^3$), particulates (31 and 291 $\mu g/m^3$), nitrogen oxides (15 and 160 $\mu g/m^3$), and hydrogen sulfide (2 and 177 $\mu g/m^3$) in the polluted city were mainly due to industrial sources. In the reference cities, air pollution was produced mainly by traffic and heating. Adjusted odds ratios (OR) for one or more upper respiratory infections of residents in the polluted city vs those in the reference cities were 2.0 (95% confidence interval [CI] = 1.3–3.2) in the younger age group and 1.6 (95% CI = 1.1–2.1) in the older age group. Within the polluted city, OR calculated for living in more vs less polluted areas were 2.0 (95% CI = 1.0–4.0) in the younger and 1.6 (95% CI = 1.0–2.7) in the older children. The present results suggests that, for children, air pollution can be hazardous in concentrations lower than those recorded in earlier studies from Britain and central Europe. The synergistic effect of sulfur dioxide, particulates, nitrogen oxides, hydrogen sulfide, and other pollutants may be a contributing factor. (*Am J Public Health.* 1991;81:1060–1063)

Introduction

Several studies in the 1960s showed that children living in areas with sulfur dioxide (SO_2) and particulate levels above 100 $\mu g/m^3$ had an increased occurrence of respiratory symptoms, upper and lower respiratory tract illnesses, and lower levels of pulmonary function than children living in less polluted areas.[1–5] The evidence is controversial on the effects of air pollution at levels lower than 100 $\mu g/m^3$.[6–9] However, a recent study from Israel reported a significantly higher prevalence of respiratory symptoms and conditions in second- and fifth-grade children living in an area of Israel with an annual average of only 30 $\mu g/m^3$ SO_2 than in children of a less polluted area.[9]

We evaluated the effect of low-level air pollution by comparing the frequency of upper respiratory infections in infants and preschool children in Oulu (polluted city), the northernmost industrial city in Finland, and in two cities with little industrial ambient air pollution, Vaasa (southern reference city) and Rovaniemi (northern reference city); see Figure 1.

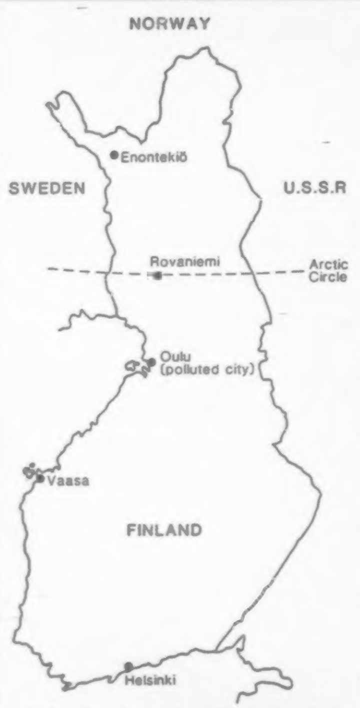

FIGURE 1—Location of Oulu, the Polluted City, and Vaasa and Rovaniemi, the Reference Cities.

Methods

Oulu (population 96 000) is located on the Baltic coast. Vaasa (coastal, population 54 000) and Rovaniemi (inland, population 31 000) are the closest southern and northern cities with an approximately similar level of urbanization but with little industrial air pollution in the downtown areas.

Four main pollution sources encircle the town center of Oulu (Figure 2). A sulfite cellulose pulp mill including a power station, northwest of the center, is the biggest source of sulfur dioxide (SO_2) pollution (2500 tonnes/y). The main source of nitrogen oxides (NO_x), a chemical plant producing acids and fertilizers, lies to the northeast. To the southwest a sulfate cellulose pulp mill, which is a source of hydrogen sulfide (H_2S) and methyl mercaptan (CH_3SH), causes odor disturbance. The heating plant of Oulu University Hospital, in the southeast, is an important source of SO_2 and particulates.

All cities monitor SO_2 and particulate concentrations. In the polluted city, H_2S and NO_x concentrations have also been measured since 1983. The measured concentrations from 1982 to 1984 are summarized in Table 1.

Jouni J. K. Jaakkola, Mikko Paunio, and Olli P. Heinonen are with the Department of Public Health, University of Helsinki. Martti Virtanen is with Jorvi Hospital, Espoo, Finland.

Requests for reprints should be sent to Jouni J. K. Jaakkola, Department of Public Health, University of Helsinki, K1/Ko 302, Otakaari 4A, SF-02150 Espoo, Finland (Fax: +358-0-451 3611).

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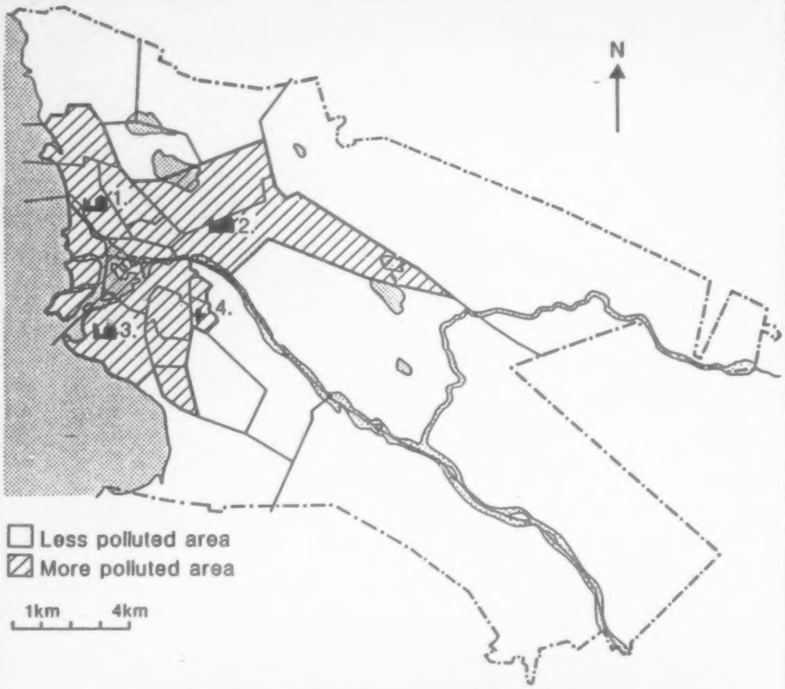

FIGURE 2—The Polluted City Divided Into Areas of More and Less Pollution: Building 1, Sulfite Cellulose Pulp Mill and a Power Plant; Building 2, Chemical Plant; Building 3, Sulfate Cellulose Pulp Mill; and Building 4, University Hospital.

Although the long-term mean levels of SO_2 were only slightly greater in the polluted city than in the reference cities, there were more frequent periods of short-term, high-level exposure to SO_2 due to downtown industrial sources. The population of the polluted city was exposed to considerably greater levels of H_2S, CH_3SH, and other malodorous sulfur compounds and nitrogen oxides.

The polluted city was divided into areas of more and less pollution according to the dispersal models of SO_2 made by the Finnish Meteorologic Institute. The more polluted area included all postal districts with an annual mean concentration of SO_2 ≥ 15 $\mu g/m^3$ and two postal districts with NO_x produced by the chemical plant but with SO_2 concentrations <15 $\mu g/m^3$. The less polluted area included those districts with an annual mean SO_2 concentration <15 $\mu g/m^3$ together with low levels of NO_x (Figure 2).

TABLE 1—Air Pollution in the Polluted City and Southern and Northern Reference Cities, 1982–1984

Pollutant ($\mu g/m^3$)	Polluted City	Southern Reference City	Northern Reference City
Sulfur dioxide			
Annual mean	13–23	9–19	17–18
Daily maximum	57–83	37–56	53
Half-hour maximum	360–807	120–190	80
Particulates			
Annual mean	31	20–33	17–28
Daily maximum	73	74–198	100–103
Half-hour maximum	291	—	—
Nitrogen oxides[a]			
Annual mean	15	—	—
Daily maximum	48	—	—
Half-hour maximum	161	—	—
Hydrogen sulfide[a]			
Annual mean	2.3	—	—
Daily maximum	42.3	—	—
Half-hour maximum	177.0	—	—

[a] Concentrations in reference cities estimated to be low due to lack of sources.

The data on health status were collected during the first two months of a mumps-measles-rubella vaccination program in November and December 1982. The total number of children ages 14 to 18 months was 679 (364, 190, and 125 from Oulu, Vaasa, and Rovaniemi, respectively) and the total of children age 6 years was 759 (310, 305, and 144, respectively).

During the vaccination visit, a nurse checked individual health records and interviewed the parents as to name, age, sex, the number of upper respiratory infections with fever (common cold, otitis media) during the previous 12 months, allergic diseases confirmed by physician (obstructive bronchitis or asthma, allergic rhinitis or conjunctivitis, and atopic dermatitis), passive smoking at home, and the number of daily contacts with other children at home and day care. The socioeconomic status of parents was obtained from the 1980 national census but was available only for older children.

The adjusted odds ratios for upper respiratory infections for those living in the more polluted vs the less polluted area were estimated in logistic regression analyses (see Appendix). The outcome was dichotomized (no infection = 0, one or more infections = 1). Similarly, other independent variables in the model were dichotomized. A child having any of the allergic diseases was considered atopic (no = 0, yes = 1). The main effects model was fitted by taking the determinant of interest (exposure) and all the other potential determinants of upper respiratory infections in the model. Meaningful interaction terms were considered according to their statistical significance ($\beta/SE[\beta] < 1.96$).

Results

In the polluted city, both younger and older children had had more upper respiratory infections during the previous 12 months than children in the reference cities (Table 2). Within the polluted city the children in the more polluted area had had more infections than children in the less polluted area. The distributions of potential confounders in the polluted and reference cities are presented in Table 3.

TABLE 2—The Occurrence of Upper Respiratory Infections in the Polluted vs Reference Cities

No. of infections during previous 12 months	Polluted city			Southern reference city n (%)	Northern reference city n (%)	Total n (%)	Unadjusted Odds Ratio[a] (95% confidence interval[b])
	More polluted area n (%)	Less polluted area n (%)	Total n (%)				
			14- to 18-month old children				
0	15 (8.0)	25 (14.2)	40 (11.0)	33 (17.4)	32 (25.6)	65 (15.5)	1.0
1–4	134 (71.3)	125 (71.0)	259 (71.1)	137 (72.1)	74 (59.2)	211 (69.2)	1.9 (1.5, 5.1)
≥5	39 (20.7)	26 (14.8)	65 (17.9)	20 (10.5)	19 (15.2)	39 (15.3)	2.7 (1.6, 4.8)
Total	188	176	364	190	125	315	
			6-year-old children				
0	29 (21.8)	53 (29.9)	82 (26.5)	92 (30.2)	67 (46.5)	159 (31.8)	1.0
1–4	97 (72.9)	118 (66.7)	215 (69.3)	202 (66.2)	76 (52.8)	278 (65.0)	1.5 (1.2, 6.0)
≥5	7 (5.3)	6 (3.4)	13 (4.2)	11 (3.6)	1 (0.7)	12 (3.2)	2.1 (0.9, 8.2)
Total	133	177	310	305	144	449	

[a] All polluted vs reference.
[b] Miettinen's test-based method.

The only difference of note was the number of daily child contacts for 6-year-old children.

Living in the polluted city vs reference cities was a significant risk factor for upper respiratory infections in both age groups when adjusting for potential confounders in logistic regression analysis (Table 4, Appendix). The adjusted odds ratio (OR) was 2.0 (95% CI [confidence interval] = 1.3–3.2) in the younger age group and 1.6 (95% CI = 1.1–2.1) in the older age group. Living in the more vs less polluted area of the polluted city was also a significant risk factor for upper respiratory infections in younger children (Table 3). The adjusted OR was 2.0 (95% CI = 1.0–4.0), i.e., a 100% increase in the frequency of children with infections. In 6-year-old children the corresponding adjusted OR was 1.6 (95% CI = 1.0-2.7), which was nonsignificant ($P < .10$), but the number of children in this analysis was small. When the less polluted area was enlarged to include surrounding municipalities, the adjusted OR became significant, 2.0 (95% CI = 1.3–3.0). There were no significant interactions in the comparisons between the more and less polluted areas.

TABLE 3—The Occurrence of Risk Factors in the Study Populations of Polluted and Reference Cities

	Polluted city			Southern reference city n (%)	Northern reference city n (%)	Total n (%)
	More polluted area n (%)	Less polluted area n (%)	Total n (%)			
			14- to 18-month old children			
Number of daily contacts						
More than 4	22 (11.7)	26 (14.8)	48 (13.7)	23 (12.1)	11 (8.8)	34 (10.8)
4 or fewer	166 (88.3)	150 (85.2)	316 (86.3)	167 (87.9)	114 (91.2)	281 (89.2)
Passive smoking at home						
Yes	57 (30.3)	46 (26.1)	103 (28.3)	51 (26.8)	38 (30.4)	89 (28.3)
No	131 (69.7)	130 (63.9)	261 (71.7)	139 (73.2)	87 (69.6)	226 (71.7)
Atopic disease						
Yes	25 (13.3)	29 (16.5)	54 (14.8)	18 (9.5)	16 (12.8)	34 (10.8)
No	163 (86.7)	147 (83.5)	310 (85.2)	172 (90.5)	109 (87.2)	121 (89.2)
			6-year-old children			
Number of daily contacts						
More than 4	111 (83.5)	134 (75.7)	245 (79.0)	280 (91.8)	107 (74.3)	387 (86.2)
4 or fewer	22 (16.5)	43 (24.3)	65 (21.0)	25 (8.2)	37 (25.7)	62 (13.8)
Passive smoking at home						
Yes	52 (39.1)	68 (38.4)	120 (38.7)	102 (33.4)	62 (43.1)	164 (36.5)
No	81 (60.9)	109 (61.6)	190 (61.3)	203 (60.6)	82 (56.9)	285 (63.5)
Atopic disease						
Yes	37 (20.9)	29 (21.8)	66 (21.3)	48 (15.7)	31 (21.5)	79 (17.6)
No	140 (79.1)	104 (78.2)	244 (78.7)	257 (84.3)	113 (78.5)	370 (82.4)

multivariate analysis. There was no municipal gas distribution in any of the cities studied. Passive smoking, daily child contacts, and atopic disease were included in the logistic models. Socioeconomic classification obtained from the 1980 census was available for the older age group. The differences in socioeconomic distribution between the cities were found to be small.

The data collection was blind: neither the participants nor the interviewers were aware of the purpose of the study. Thus the potential information bias caused by knowledge of children's exposure status was eliminated.

In conclusion, our results suggest that children's respiratory health can be jeopardized from exposure to air pollution in concentrations much lower than those recorded in earlier studies from Britain and central Europe. The synergistic effect of SO_2, particulates, NO_x, H_2S, and other pollutants may be a contributing factor. ☐

TABLE 4—Crude (cOR) and Adjusted Odds Ratios (aOR) for Having Experienced an Upper Respiratory Infection during the Past 12 Months in Polluted vs Reference Cities and in the More vs Less Polluted Area of the Polluted City

Population	cOR	aOR	95% CI[a]
14- to 18-month-old children			
Reference cities	1.0	1.0	
Polluted city	2.1	2.0	1.3—3.2
Less polluted area	1.0	1.0	
More polluted area	1.9	2.0	1.0—4.0
6-year-old children			
Reference cities	1.0	1.0	
Polluted city	1.5	1.6	1.1—2.1
Less polluted area	1.0	1.0	
More polluted area	1.5	1.6	1.0—2.7

Adjustment was made by logistic regression analysis controlling for sex, daily child contacts, atopic disease, and passive smoking.

[a] CI = confidence interval.

Discussion

The ambient air in the index city was more polluted than in the reference cities due to the industrial sources around the town center. There were also differences in air pollution across the polluted city. However, the differences in the annual mean concentrations of SO_2 between the cities and within the polluted city were small. Moreover, in terms of international comparison, the levels in the most polluted area in the polluted city (25 $\mu g/m^3$) can still be considered low-level air pollution. However, the industrial sources in the polluted city cause higher and more frequent peaks of SO_2 than are caused by heating and traffic in the reference cities. Industrial NO_x, H_2S, and CH_3SH emissions were the main qualitative differences in comparing the polluted city with the reference cities.

We found a significant association between the occurrence of upper respiratory infections and living in an air-polluted area. This finding was consistent in both 14 to 18 month olds and 6 year olds (and combined) when comparing the polluted city with the reference cities and when comparing the more and less polluted areas within the polluted city. Most of the potential confounders were controlled for either by the study design and/or in the

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APPENDIX—Regression Coefficients Produced by PROBIT Procedure of SPSSX

Risk factor	Polluted vs reference city: Younger children β (SE)	Polluted vs reference city: Older children β (SE)	Polluted vs reference area[a]: Younger children β (SE)	Polluted vs reference area[a]: Older children β (SE)
Living in polluted city or area	0.357 (0.110)	0.221 (0.082)	0.339 (0.174)	0.256 (0.139)
More than 4 daily child contacts	0.423 (0.306)	0.121 (0.099)	0.476 (0.522)	−0.289 (0.161)
Atopic disease	0.224 (0.187)	0.140 (0.105)	0.152 (0.254)	0.073 (0.163)
Passive smoking	0.114 (0.124)	−0.140 (0.081)	−0.058 (0.185)	−0.084 (0.133)
Male sex	−0.131 (0.108)	0.025 (0.079)	−0.115 (0.170)	0.109 (0.131)
Intercept	5.674	5.217	5.926	5.596

Note: All the variables are coded yes = 1 and no = 0.

The final equation for polluted vs reference city, younger children is:

logit [pr(INFECTION = 1 /POLLUTED, CONTACTS, ATOPIC, PASSIVE, SEX)] = Intercept + 0.357(POLLUTED) + 0.423(CONTACTS) + 0.224 (ATOPIC) + 0.114(PASSIVE) − 0.131(SEX)

The procedure calculates the estimates from LOGIT model:

$$\{LOG[p/(1-p)]\}/2 + 5 = \text{Intercept} + \beta X$$

This means that aOR can be calculated:

$$aOR = e^{2\beta}$$

and the 95% confidence limit is

$$e^{2\beta \pm [illegible]}$$

[a] Restricted to the residents of the polluted city.

Decline in the Use of Illicit Drugs by High School Students in New York State: A Comparison with National Data

Denise B. Kandel, PhD, and Mark Davies, MPH

ABSTRACT

A school-based epidemiological survey of the use of drugs was carried out in Spring 1988 among 7,611 adolescents representative of students in grades 7 through 12 in New York State public and private schools. This survey provides the most recent data on the use of crack in that population. Comparison with a statewide survey carried out five years earlier reveals a large decline in the use of illicit drugs in that period. This decline appears to be stronger than the decline reported for national samples. (*Am J Public Health.* 1991;81:1064–1067)

Introduction

National epidemiological surveys[1,2] document a decrease in the use of illicit drugs starting in 1980 for marijuana and in the mid-1980s for cocaine. We report results from a 1988 survey of high school students in New York State (NYS) which reveals large declines in the use of illicit drugs over five years. The decline appears to be stronger in the Northeast than in the United States as a whole and sharper in NYS than in the entire Northeast.

Methods

A statewide epidemiological survey of the use of alcoholic beverages, cigarettes, marijuana, cocaine, crack, and other illicit drugs was conducted among 7,611 students in grades 7 through 12 in NYS in Spring 1988.

The two-stage random sample represents junior and senior high school students attending NYS public and private schools. A stratified sample of 54 schools and two homerooms from each grade per school was selected. The four stratification criteria for school selection were: 1) geographical area; 2) proportion White enrollment; 3) public versus private; 4) size of enrollment. Active or passive parental consent for the child's participation was obtained, as required by individual schools. A school in New York City was dropped because only 10 percent of parental consent forms were returned. Students answered anonymous self-administered structured questionnaires in classrooms (84 percent completion rate). The sample was weighted to reflect the variable probabilities of selection of schools and homerooms and the grade-specific absentee rate in each school; 95 percent confidence intervals were calculated using empirical standard errors from 10 random subsamples drawn from the total sample. For grade-specific rates, the standard error was estimated by multiplying the standard error, assuming simple random sampling, by the average design effect (1.53).

Results

Prevalence of Student Drug Use

The rates of use follow patterns reported by others,[4,5] but probably underestimate drug use among adolescents in the population, since school absentees and dropouts were not included (Table 1).[6–9]

Trends Over Time

In 1983, the NYS Division of Substance Abuse Services (DSAS) surveyed a representative sample of students in grades 7 to 12 from 206 public and private schools selected from a stratified sample of health service areas.[3] Structured self-administered questionnaires were given in classrooms. (Completion rates are not provided.).

Comparison of rates in 1983 and 1988 reveals sizable declines of at least 50 percent over five years in the use of almost every illicit drug, but no change in alcohol use (Table 1).

Comparison with National Data

The declines in NYS schools are striking and parallel those reported nationally for high school seniors in the Monitoring the Future (MF) study by Johnston, *et al.*[1] and for adolescents ages 12 to 17 in the population reported in the National Household Survey (NHS)[2] (Table 2). [Lower rates are reported in the household (NHS) than the school (NYS) survey,[10] and may reflect in part differences in mode of survey administration.] This decline may be greater in NYS and the Northeast than nationally. Over two years, the prevalence of cocaine use reported by high school seniors declined by 43 percent in the Northeast (23 percent in

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Address reprint requests to Denise B. Kandel, PhD, Department of Psychiatry, School of Public Health, Columbia University, 722 West 168th Street, Box 20, New York, NY 10032. Mr. Davies is also with that Department; Dr. Kandel is also affiliated with New York State Psychiatric Institute. This paper, submitted to the Journal June 11, 1990, was revised and accepted for publication December 7, 1990.

TABLE 1—Lifetime Prevalence Rates of the Use of Various Drugs among 7th through 12th Grade Students in New York State in 1988 and 1983 and by Grade in 1988 (95% Confidence Intervals)

	Alcohol % (CI)	Cigarettes % (CI)	Marijuana % (CI)	Stimulants % (CI)	Inhalants % (CI)	Any Cocaine % (CI)	Crack % (CI)	Psychedelics % (CI)	Sedatives % (CI)	Tranquilizers % (CI)	Heroin % (CI)	N
All 7th–12th Graders												
NYS 1988	84.4 (1.5)	57.7 (2.2)	27.5 (1.3)	11.1 (0.8)	6.3 (0.6)	5.9 (0.5)	1.9 (0.4)	4.8 (0.7)	4.1 (0.7)	3.6 (0.4)	1.4 (0.4)	(7,611)
NYS 1983[a]	83	N/A	46	25	22	14	N/A	10	12	12	3	(27,414)
Grade-Specific NYS 1988												
7th grade	66.1 (4.2)	36.0 (4.2)	3.7 (1.7)	1.7 (1.1)	2.8 (1.5)	1.5 (1.1)	.8 (0.8)	1.1 (0.9)	.9 (0.8)	1.0 (0.9)	1.1 (0.9)	(1,210)
8th grade	78.2 (3.7)	51.9 (4.5)	12.2 (3.0)	2.8 (1.5)	4.4 (1.9)	1.8 (1.2)	.7 (0.8)	2.2 (1.3)	1.6 (1.1)	2.0 (1.3)	.6 (0.7)	(1,149)
9th grade	84.5 (3.0)	58.6 (4.0)	24.3 (3.5)	11.2 (2.6)	6.0 (2.0)	4.2 (1.6)	2.3 (1.2)	4.6 (1.7)	4.4 (1.7)	4.8 (1.8)	2.2 (1.2)	(1,367)
10th grade	88.7 (2.6)	61.1 (4.0)	30.1 (3.8)	14.1 (2.9)	6.7 (2.1)	5.4 (1.9)	1.9 (1.1)	5.4 (1.9)	4.7 (1.8)	3.8 (1.6)	1.1 (0.9)	(1,361)
11th grade	93.0 (2.1)	66.4 (3.9)	42.8 (4.1)	16.8 (3.1)	8.1 (2.3)	8.5 (2.3)	2.1 (1.2)	5.7 (1.9)	5.6 (1.9)	4.2 (1.7)	1.6 (1.0)	(1,348)
12th grade	94.7 (2.0)	70.9 (4.1)	50.9 (4.5)	18.8 (3.5)	9.7 (2.7)	14.0 (3.1)	3.3 (1.6)	9.6 (2.7)	7.4 (2.4)	5.5 (2.1)	1.8 (1.2)	(1,156)

N/A = Not ascertained. Cigarettes use ascertained for last month not lifetime.
[a]Source: DSAS, 1984.[9] Confidence intervals not available.

1986 versus 13 percent in 1988[1,11]), compared to 28 percent nationally (17 percent versus 12 percent). We compared the changes in rates in NYS over comparable periods in MF (1983–1988) and NHS (1982–88) (Table 3). For NYS, we calculated odds ratios for 11th–12th graders, since separate 1983 data for the 12th grade were not available.

The odds ratios are consistently and significantly higher in NYS than in the nation, except for sedatives and tranquilizers. In MF, the ratios are slightly higher in the Northeast than in the total sample and significantly different for marijuana, cocaine and stimulants.

Discussion and Conclusion

The decline in drug use appears to be sharper in areas where rates were initially the highest. This differential decline may be explained by changes in the age structure of the US population, peer influence processes, and the relative commitment of youth to substance use in areas with high compared to low use prevalence. Substance use is most prevalent in the late teens and early twenties.[2,5] The downward trend in illicit drug use prevalence from 1980 to 1988 parallels the decline in the ratio of youths (aged 15–24) to the parental generation (aged 34–44) between

TABLE 2—Comparisons of Prevalence Rates of Lifetime Drug Use among High School Seniors in Monitoring the Future in 1988 and 1983 Nationally and in the Northeast and among 12- to 17-Year-Olds in 1988 and 1982 in the National Household Survey

	Alcohol % (CI)	Cigarettes % (CI)	Marijuana % (CI)	Stimulants % (CI)	Inhalants % (CI)	Any Cocaine % (CI)	Crack % (CI)	Psychedelics % (CI)	Sedatives % (CI)	Tranquilizers % (CI)	Heroin % (CI)	N
Monitoring the Future												
12th Grade												
88 National[a]	92.0 (1.5)	66.4 (1.7)	47.2 (2.1)	19.8 (1.5)	16.7 (1.3)	12.1 (1.2)	4.8 (0.7)	8.9 (1.0)	7.8 (1.0)	9.4 (1.1)	1.1 (0.3)	(16,796)
83 National[b]	92.6 (1.3)	70.6 (1.5)	57.0 (2.1)	26.9 (1.5)	18.2 (1.2)	16.2 (1.4)	N/A	11.9 (1.0)	14.4 (1.3)	13.3 (1.3)	1.2 (0.3)	(16,300)
88 Northeast[a]	93.9	66.6	49.6	16.5	15.3	13.2	3.8	9.3	7.7	8.1	1.0	(3,345)
83 Northeast[b]	95.4	72.9	63.7	26.9	13.0	20.5	N/A	14.0	12.4	12.3	1.1	(4,056)
National Household Survey												
12–17 year olds												
88 National[d]	50.2 (5.7)	42.3 (4.2)	17.4 (4.1)	4.2 (2.1)	8.8 (2.1)	3.4 (1.7)	.9 (0.8)	3.5 (1.5)	2.4 (1.8)	2.0 (1.5)	.6 (0.5)	(3,095)
82 National[c]	65.2 (3.0)	49.5 (3.1)	26.7 (2.7)	6.7 (1.6)	N/A	6.5 (1.6)	N/A	5.2 (1.4)	5.8 (1.5)	4.9 (1.4)	<.5	(1,581)
88 Northeast[d]			17.3									(566)
82 Northeast[c]			31.0									(337)

N/A = Not ascertained.
Sources: [a]Monitoring the Future 1988[1]—12th graders. Standard errors not available for Northeast.
[b]Monitoring the Future 1983[10]—12th graders. Standard errors not available for Northeast.
[c]National Household Survey 1982[11]—12–17 year olds. Northeast data not presented for age specific groups, except marijuana; standard errors not available.
[d]National Household Survey 1988[2]—12–17 year olds. Northeast data not presented for age specific groups, except marijuana; standard errors not available.

TABLE 3—Odds Ratios for Changes in Rates of Lifetime Drug Use Over a 5-Year Period among Adolescents in National Samples, the Northeast and New York State from 1983 (or 1982) to 1988 Odds Ratios

	Any Alcohol % (S.E.)	Cigarettes % (S.E.)	Marijuana % (S.E.)	Stimulants % (S.E.)	Inhalants % (S.E.)	Any Cocaine % (S.E.)	Psychedelics % (S.E.)	Sedatives % (S.E.)	Tranquilizers % (S.E.)
Monitoring the Future[a]									
12th graders (1983–1988)									
National	1.1 (.05)	1.2 (.03)	1.5 (.03)	1.5 (.09)	0.8 (.02)	1.4 (.04)	1.4 (.05)	2.0 (.07)	1.5 (.06)
Northeast	1.4 (.15)	1.4 (.07)	1.8 (.09)	1.9 (.11)	0.8 (.05)	1.7 (.11)	1.6 (.12)	1.7 (.14)	1.3 (.11)
National Household Survey[b]									
(12–17 years old) (1982–1988)									
National	1.9 (.12)	1.3 (.08)	1.7 (.13)	1.6 (.22)	N/A	2.0 (.28)	1.5 (.22)	2.5 (.40)	2.5 (.43)
Northeast	N/A	N/A	2.2 (.36)	N/A	N/A	N/A	N/A	N/A	N/A
New York State[c] (1983–1988)									
7th–12th graders	.9 (.03)	N/A	2.2 (.06)	2.7 (.11)	4.2 (.21)	2.6 (.13)	2.2 (.13)	3.2 (.19)	3.6 (.23)
11th–12th graders	N/A	N/A	2.1 (.10)	2.7 (.15)	2.3 (.18)	2.1 (.15)	2.3 (.19)	2.2 (.19)	1.2 (.12)

N/A = Data not available in published reports.

Sources: [a]Johnston, O'Malley, and Bachman.[10,1]
[b]Miller et al,[11]; NIDA.[2]
[c]DSAS[3]; present study.

the early and late 1980s, from 1.7 to 1.1. (The upward trend in prevalence from 1960 to 1980 paralled the upward trend in the ratio of youths to adults.) Fewer members in one's age cohort and smaller relative cohort size will reduce opportunities for social interactions with one's peers, a most important factor in drug use initiation,[14] and increase social control by the older generation.[15,16] Perceived risks and disapproval associated with illicit drug use have increased.[17,18] Although changes in relative cohort sizes are almost identical in NYS and the US, the sharper decline in drug use prevalence in NYS could be accounted for by the structural factor, if changes in individual behavior depend not only on opportunities for peer interactions but on group norms and drug users' characteristics. The rate of decline in drug use may follow a reverse diffusion process characteristic of epidemics, and may accelerate as a function of the number of individuals initially exposed, especially if proscriptive norms also become more negative and individuals at risk for drug involvement are less deviant. In areas with high prevalence of drug use, less deviant youths will be drawn into drug use, mainly through peer influence. The high use areas would show more rapid decline in prevalence than areas with low prevalence, where the pool of existing and potential users would include individuals more committed to drugs.

The decrease in lifetime drug use observed among NYS students cannot be attributed to an increase in school dropout rates by drug-using adolescents, thereby leaving in schools an increasing number of non-drug users, since the dropout rate has been declining.[15]

These epidemiological data contrast with data on treated or clinical cases, which have shown dramatic increases throughout the 1980s,[20] although recent data on cocaine-related emergency room admissions show declines.[21] There may be a time-lag between initiating drug use and experiencing drug-related problems; smoked cocaine now accounts for a large part of clinical cases.

The most extreme forms of drug use may be concentrated in disorganized communities that are not well represented in school or household samples. Whether the situation in these communities has deteriorated over the last decade is not known. We may be moving toward an increasingly polarized society, in which drug involvement is another dimension, besides lower economic and social resources and higher social pathologies,[22] on which groups are divided and the gap between them is becoming greater.[22,23] There would be youths in school and mainstream society, who have decreased their drug use, and those in urban ghettos and out of school, who have increased their use.[24] Strengthening the educational system and youth's investment in their schooling and future may be a most promising way of reducing involvement in such activities as drug use, which are destructive for self and society. ☐

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Surveillance of Attempted Suicide among Adolescents in Oregon, 1988

Jon K. Andrus, MD, David W. Fleming, MD, Michael A. Heumann, MPH, James T. Wassell, PhD, David D. Hopkins, MS, and Jane Gordon, PhD

ABSTRACT

In January 1988, Oregon became the first state to require hospital-based reporting of attempted suicide (AS) in all adolescents < 18 years old. From January to December 1988, 644 cases of AS were reported (annual rate of 214 per 100 000 population, ages 10 to 17 years). We compared these 644 cases of AS with all 137 Oregon adolescents < 18 years old who committed suicide in Oregon during the 10-year period 1979 through 1988, and found that the strongest predictor of outcome was method used. (*Am J Public Health.* 1991;81:1067–1069)

Introduction

Nationally, suicide is the third leading cause of death among adolescents and young adults.[1] For each completed suicide, 30 to 200 suicides are attempted, but not completed.[2–6] Ten percent of adolescents who attempt suicide make further attempts within 1 year, and 31% do so within 2 years.[7,8]

In Oregon, suicide among teenagers is second only to motor vehicle injuries as a cause of death, accounting for 16% of deaths in Oregonians ages 13 to 19.[9,10] In 1987 the Oregon legislature mandated the reporting of all attempted suicides among persons younger than 18 years. This paper summarizes the first year (1988) of information from this surveillance system and compares the characteristics of adolescents who attempted suicide in 1988 with those who completed suicide in Oregon during the 10-year period 1979 to 1988.

Methods

Beginning January 1988, each Oregon hospital (n = 70) was required to report to the Oregon Health Division information on all adolescents younger than 18 years who attempted suicide. An attempted suicide was defined as self-inflicted injury or condition specified by the medical provider as having fatal intent, treated at a hospital or a hospital emergency department.

Information collected by the hospital included demographics, date and place of attempt, family living situation, history of previous suicide attempts, use of alcohol, results of blood alcohol testing if done, and method of attempt. This information, without name identifiers, was sent monthly to the Health Division.

For comparison, state death certificates and medical examiner's reports were used to identify all Oregon adoles-

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Jon K. Andrus and James T. Wassell are with the Epidemiology Program Office of the Centers for Disease Control, Atlanta, Ga. David W. Fleming, Michael A. Heumann, David D. Hopkins, and Jane Gordon are with the Oregon Health Division, Office of Health Status Monitoring, Portland, Ore.

Requests for reprints should be sent to Jon K. Andrus, MD, Pan American Health Organization, HPM/EPI 7th Floor, Room 714, 525 23rd Street, NW, Washington, DC 20037.

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TABLE 1—Gender, Race, and Place of Residence of Attempted and Completed Suicide[a] among Adolescents in Oregon

Characteristic	Attempters No.	Attempters %	Completers No	Completers %	OR[b]	CI[c]
Gender						
Female	538	83.6	27	19.7	20.8	12.7–34.3
Male	106	16.4	110	80.3		
Race						
White	547	84.9	127	92.6	2.2	1.1–4.9
Other	97	15.1	10	7.4		
Residence						
Urban	128	19.8	25	18.2		
Suburban	256	39.8	47	34.3		
Rural	260	40.3	65	47.4		

[a] For attempters annual rate for 1988; for completers, average annual rate for 1979–1988.
[b] OR = odds ratio.
[c] CI = confidence interval.

TABLE 2—Age-Specific Annual Incidence Rates of Attempted and Completed Suicide among Adolescents in Oregon[a]

Age (years)	Attempters No[b]	Attempters Rate[c]	Completers No.	Completers Rate[c]	Rate ratio (A/C)[d]
10–11	6	7.2	4	0.5	14.4
12–13	72	91.6	9	1.1	83.2
14–15	248	298.3	38	4.5	66.3
16–17	309	340.6	86	9.5	35.9
Total	635	214.0	137	4.6	46.5

[a] For attempters, annual rate for 1988; for completers, average annual rate for 1979 to 1988.
[b] Age is unknown for nine attempters.
[c] Rate per 100 000 age-specific population.
[d] A = attempters; C = completers.

TABLE 3—Method Used for Attempted and Completed Suicide[a] among Adolescents in Oregon

Method	Attempters No.	Attempters %	Completers No.	Completers %
Drug overdose	486	75.5	6	4.4
Laceration or stabbing	67	10.4	0	0.0
Poisons or caustics	42	6.5	1	0.7
Hanging	10	1.6	23	16.8
Firearms	4	0.6	99	72.3
Other or unknown[b]	35	5.4	8	5.8
Total	644	100.0	137	100.0

[a] For attempters, annual rate for 1988; for completers, average annual rate for 1979–1988.
[b] Other or unknown methods include jumping, drowning, motor vehicle crash, unknown.

cents 10 to 17 years old who committed suicide in Oregon during the years 1979 to 1988. A completed suicide was defined as a person who died as a result of intentional self-inflicted injuries as determined by the medical examiner.

In this study, the term adolescent means 10- to 17-year-olds. Adolescents who attempted suicide but were not successful are termed attempters, adolescents who committed suicide are termed completers, and incidence rates are expressed per 100 000 population of 10- to 17-year-olds. Rural counties were those with less than 15 000 population; suburban were those with populations of 15 000 to 35 000; and urban were those with greater than 35 000 population.

Population data were obtained from the Bureau of the Census and the Center for Population Research and Census at Portland State University. For attempters, we used 1988 mid-year population estimates for denominators to calculate annual incidence rates. For completers, we used 1984 population estimates. Annual incidence rates were calculated for completers using 10 years of data from 1979 to 1988. This methodology was chosen after we had compared the characteristics of completers from 1979 to 1983 with those from 1984 to 1988 and found no differences by age, sex, race, or method of attempt.

Data were analyzed using Dbase, EPIINFO, and SAS. Exact confidence limits were calculated using methods based on Fisher's Exact Test.

Results

In 1988, 644 Oregon adolescents ages 10 to 17 years were reported to have attempted suicide, an annual incidence rate of 214 per 100 000. There were no attempted suicides in children younger than 10 years. Of attempters, 87.7% were aged 14 to 17 years, 83.5% were female, 84.9% were White, and 19.8% were urban residents (Table 1).

Only 9.5% (61/644) of the attempters had blood alcohol tested. Of those tested, 50.8% had detectable blood alcohol levels. The medical examiner tested 41.6% (57/137) of the completers for blood alcohol. Of completers tested, 35.1% had documented detectable blood alcohol at the time of their death.

From 1979 to 1988, 137 Oregon adolescents committed suicide, an average annual incidence rate of 4.6 per 100 000. The ratio of the average annual incidence rates of attempters to completers was 47:1. The average annual incidence rates for both attempted and completed suicide was lowest for 10 to 11 year olds and highest for 16 to 17 year olds (Table 2).

Compared with attempters, completers were more than 20 times more likely to be male and 2 times more likely to be White (Table 1). There were no statistically significant differences between attempters and completers by residence, place of attempt, day of the week, or season when the self-inflicted injury occurred.

Attempters more frequently used drug overdose as a method, while completers more frequently used firearms (Table 3). Gunshot wounds and hanging were strong predictors of fatal outcome. These two methods were used by 89.1% (122/137) of the completers compared with 2.2% (14/644) of the attempters (odds ratio [OR] = 366.0, 95% confidence interval [CI] = 162.8–844.8).

We performed a logistic regression analysis including the method, sex, race, season, residence, day of the week, and age. Choice of method was the strongest predictor of outcome ($P < .0001$). Although sex had the only other statistically significant predictor coefficient ($P = .002$), it had little effect on the predicted probability of outcome.

The univariate association of sex with outcome was primarily due to an association of sex with method. Combining data from attempters and completers showed that lethal methods were more likely to be chosen by males: 90.9% (30/33) of the hangings and 79.6% (82/103) of the shootings. In contrast, among those using stabbing or lacerations, 20.9% (14/67) were male; and among the drug overdosers, 13.0% (64/492) were male.

Discussion

In 1988, one adolescent suicide attempt with injuries severe enough to require emergency care was reported for every 467 Oregon adolescents. The ratio of attempted to completed suicide rates for those adolescents was 47:1. The overwhelming predictor of outcome was method used. Generally, adolescents who shot themselves died, while those who overdosed on drugs or lacerated themselves lived.

Collection of accurate prevalence or incidence data for attempted suicide has been hampered by a lack of central registries.[6,11] Results of limited surveys of high school students vary and have estimated that 5% to 15% of adolescents attempt suicide.[12–15] Our study showed that adolescent suicide attempts are common in Oregon and that statewide surveillance is feasible.

In this study, method and outcome were found to be inextricably linked. Thus the factors influencing one must also influence the other. Whether desired outcome determines method or the method available at the time of attempt determines outcome urgently needs clarification. For some adolescents who attempt suicide, selection of method may be paramount, and outcome will depend on method. For others, desired outcome may be decided first and selection of method will depend on that decision. Successful intervention is critically dependent on which of these two dynamics is operating. Securing firearms in a locked cabinet may not be helpful if a despondent adolescent will merely search for another lethal weapon. However, this intervention would be quite effective if the result were that an alternate, less lethal method would be used. Additional study is needed to determine factors influencing choice of method and to identify situations in which selection of method is not solely dependent on desired outcome, but is also influenced by other factors such as availability and social acceptability. □

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Nurses Claim Victory in Supreme Court Decision

"Registered nurses across the country have won a major victory in the United States Supreme Court," Dr. Barbara Redman, Executive Director of the American Nurses Association (ANA), recently announced.

The court upheld an ANA-backed position allowing for separate collective bargaining units in hospitals for registered nurses.

"This decision gives nurses the right to decide their destiny in the workplace," said Redman. "Nurses in many states across the nation have told us they are ready now to organize, and the ANA has allocated half a million dollars to state nurses associations to train organizers and help increase their organizing efforts."

"The ANA is uniquely positioned to handle an increased demand by nurses for assistance in determining workplace issues. We are the only national organization offering collective bargaining that is run by nurses for nurses," added Redman.

Redman said the ruling will lead to a stronger voice for nurses in promoting higher quality care for patients, adding value to the health care dollar.

Knowledge of Diagnosis and Reproductive History among Survivors of Childhood Plumbism

Howard Hu, MD, ScD, MPH

ABSTRACT

As part of a study of the long-term consequences of childhood plumbism, 35 survivors of lead poisoning in the period 1930 through 1944 were interviewed along with 22 controls matched for age, sex, and town of residence. Subjects were asked to recall all pregnancies they engendered (men) or carried (women), outcome, and intellectual development of resulting children. Matched female plumbism subjects reported a higher proportion of spontaneous abortion or stillbirths among pregnancies (relative risk = 1.60, 95% confidence interval [CI] = 0.6–4.0) and a higher proportion of learning disabilities among school-age children (relative risk = 3.0, 95% CI = 0.9–10.2) in comparison with their controls. There was no evidence of clustering of learning disabilities within families or confounding by parental educational level. These results suggest that women with a history of childhood lead poisoning may be at risk for having spontaneous abortions or stillbirths and having children who manifest significant learning disabilities. (*Am J Public Health.* 1991;81:1070–1072)

Introduction

Lead accumulated from past exposures is not inert, as previously thought; rather, it is probably released continually from storage sites.[1,2] More than 95% of stored lead in an adult is located in the skeleton. Recent mammalian animal studies have suggested that a significant fraction of skeletal lead content is mobilized, enters the circulation, and crosses the placenta during both pregnancy and lactation.[3,4]

Since lead is a well-known behavioral teratogen, even at relatively low levels of exposure, and skeletal lead stores are known to persist for decades, the possibility exists that lead can be a delayed transgenerational toxin, i.e., it can affect the next generation even after maternal exposures have long ceased.

We report results of a reproductive questionnaire administered to subjects participating in an investigation of the long-term consequences of childhood plumbism (lead poisoning). We also report the level of awareness that survivors of childhood plumbism had with regard to personal histories of lead exposure. All research described herein was approved by the institutional review boards of the Brigham and Women's Hospital, Boston Children's Hospital and Medical Center, and the Harvard School of Public Health.

Methods

Methods used to define, trace, and recruit eligible subjects are described in detail elsewhere.[5] In summary, we used microfilmed medical records to identify 192 subjects admitted to Boston Children's Hospital during the period 1930 to 1944 who met the following criteria for lead poisoning: (1) a definite history of pica for lead-containing material (usually paint) or the use of lead nipple shields, (2) evidence of dense metaphyseal bands ("lead lines") on at least one long bone x-ray, (3) clinical evidence of symptoms and signs characteristic of childhood lead intoxication, and (4) the absence of clinical evidence for rickets. Subjects with continued residence in Massachusetts were traced using available public records.

The current status of 145 (76%) subjects was ascertained. Of these, 102 were found with current Massachusetts addresses. The 72 who lived in the Boston area were sent letters recruiting them for a "study on the impact of the environment on health."

Thirty-five subjects participated in the study. In general, participants had a history of less severe plumbism than nonparticipants.[5] Also enlisted were 22 control subjects matched for age, sex, and neighborhood. Town books were used to identify eligible controls, and letters of recruitment were sent. Budget constraints prevented recruitment of matched control subjects for 13 of the childhood plumbism subjects.

Each participant received a package containing a self-administered questionnaire and instructions. The questionnaire included detailed medical, occupational, and environmental history components. Each subject was asked to list all known pregnancies (or engendered pregnancies, in the case of males) with dates; whether the pregnancy resulted in a miscarriage, still birth, live birth, or was voluntarily aborted; and the sex of live births. For each offspring that survived to school age, the subject was asked whether he or she had had a learning disability. This was defined as difficulty in school that either (1) resulted in being left back one or more grades or (2) required special in-school remedial education. Learning disability de-

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Howard Hu is with Channing Laboratory, Brigham and Women's Hospital, Harvard Medical School, Boston and the Occupational Health Program, Dept. Environmental Health, Harvard School of Public Health, Boston. This research was conducted while the author was with the Department of Epidemiology, Harvard School of Public Health, Boston.

Requests for reprints should be sent to Dr. Howard Hu, Channing Laboratory, Brigham and Women's Hospital, Harvard Medical School, 180 Longwood Avenue, Boston, MA 02115.

This paper was submitted to the journal July 2, 1990, and accepted with revisions February 20, 1991.

fined similarly has been found to be associated with elevated dentine levels in a longitudinal study of teenagers.[6]

Results

Matched plumbism subjects did not differ appreciably from either the matched controls or the nonmatched plumbism subjects with respect to age, sex ratio, years of education, smoking status, alcohol ingestion, and medication use (Table 1).

In response to questions on childhood illnesses, 13 (37%) of all plumbism subjects noted being unaware of having had lead poisoning as children.

Among all plumbism subjects, women had 19 miscarriages or stillbirths out of a total of 79 pregnancy outcomes (24%), and 13 children with learning disabilities out of 60 school-age children (22%) (Table 2). Matched female plumbism subjects had 11 miscarriages or stillbirths out of a total of 51 pregnancies (22%), and 10 children with learning disabilities among a total of 40 school-age children (25%). Matched control subjects had 6 miscarriages or stillbirths out of 48 pregnancy outcomes (13%), and 3 learning-disabled children out of 42 school-age children (7%). The relative risk of matched female plumbism subjects for having a miscarriage or stillbirth was 1.60 (95% confidence interval [CI] = 0.6–4.0) and for having a learning disabled child, 3.0 (95% CI = 0.9–10.2). Inclusion of unmatched plumbism subjects did not appreciably alter these estimates. The pregnancies engendered by both the matched plumbism and control male subjects had few adverse reproductive outcomes. There were too few matched pair outcomes to offer meaningful analysis using McNemar's test.

TABLE 1—Participant Demographics and Severity Scores

Characteristic	Matched Plumbism Subjects	Matched Control Subjects	Unmatched Plumbism Subjects
n	22	22	13
Age			
Range (years)	49–60	48–60	47–60
Mean (years) ± SD	54.7 ± 4.0	55.4 ± 4.0	56.7 ± 4.0
Ratio of males to females	8/14	8/14	5/8
Education			
Range (years)	8–20	8–18	11–16
Mean (years) ± SD	13.5 ± 2.8	14.3 ± 2.9	12.9 ± 1.9
Smoking			
Never smokers	8	7	5
Former smokers	11	12	5
Current smokers	3	3	3
Usual alcohol ingestion			
<2 Alcoholic drinks per day	19	20	12
≥2 Alcoholic drinks per day	3	2	1

Discussion

The developing fetus and infant is clearly susceptible to the neurobehavioral effects of lead exposure. Efforts are currently under way to decrease such susceptibility by decreasing maternal exposures that have been associated with elevated lead levels in the blood, such as the use of leaded gasoline,[7,8] smoking and alcohol use,[9] and occupational exposures.[10]

Although this inquiry is of a small scale and produced results of borderline significance defined statistically, it provides evidence that remote exposures such as childhood plumbism may cause subsequent risk to a fetus or infant and result in later impairment of school performance. The relative risks found in this study may be underestimates, since study participants had been less severely intoxicated than nonparticipants.

This research also suggests that a significant fraction of individuals who had been diagnosed with lead toxicity as children are unaware of this history.

The presence of significant bias in this study seems unlikely. Other than a childhood history of plumbism, matched subjects did not differ from controls in terms of age, years of education, medication use, smoking, or alcohol ingestion. Studies that have examined the consistency of reproductive history questionnaires have generally not found significant recall bias.[11] Our analysis did not account for familial aggregation of adverse outcomes.[12] However, no more than two learning disabilities were observed in any one family in this study.

These results justify additional research efforts in this area. Overexposure to lead is still a widespread phenomenon among our society's children.[13] If lead poisoning poses a risk not only to these children, but also to these children's children, it may be important to screen young women entering reproductive age for lead burden and develop a strategy for prevent-

TABLE 2—Summary of Reponses to Reproductive Questionnaire

	Pregnancies No.	Miscarriages or Stillbirths No.	Percentage	Children Alive at School Age	Children with Learning Disabilities No.	Percentage
Females						
Matched plumbism subjects	51	11	22	40	10	25
Matched control subjects	48	6	13	42	3	7
Nonmatched plumbism subjects	28	8	29	20	3	15
Males						
Matched plumbism subjects	30	1	3	29	3	10
Matched control subjects	30	3	10	27	2	7
Nonmatched plumbism subjects	16	0	0	16	1	6

ing mobilization and/or transfer of accumulated lead to the fetus. History-taking may not be sufficient to identify women at risk. □

Acknowledgments

This research was supported by National Institutes for Environmental Health Sciences Center (NIEHS) Grant ES00002, NIEHS National Research Service Award 2T 32ES 07069, and National Institute of Health Clinical Research Center Grant RR 02635. The project would not have been possible without the diligent tracing efforts of Nancy Potter and her assistants at the Dana Farber Cancer Center. Many thanks go to Drs. Richard Monson, Stephen Lagakos, David Christiani, John Graef, and Fred Li for expert advice; Sarah Freeman, Kim Damakosh, and Cynthia Cryan for help in recruiting traced subjects; and Susan Fisher, RN, in evaluating subjects.

An abstract of this research was presented at the August 1990 meeting of the International Society for Environmental Epidemiology in Berkeley, CA.

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Editor Search Committee Seeks New Editor for American Journal of Public Health

An Editor Search Committee is actively seeking a replacement for Dr. Michel Ibrahim, who has announced his resignation as Editor of this Journal, effective January 1, 1992. The six-member search committee welcomes and solicits nominations and recommendations from the APHA membership and leadership to assist them in identifying the best available individual for this important salaried position. Nominations and statements of interest should be submitted by August 20, 1991; the committee will meet several times to consider suggestions, applications, and support materials of potential candidates.

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spect to the AIDS epidemic: (1) Could an immediately available, nonpunitive maintenance program attract into treatment a significant number of presently unmotivated addicts and reduce their risk behavior? (2) If a large number responded, could clinics of this kind expand fast enough to meet demand, while keeping a balance with other programs providing full service? The authors approached these questions by restricting the study group to addicts on the waiting list of an established full-service program, with the understanding that they would be released from the study when an opening became available in the conventional program.

Even with this reduction in scope, the study was vigorously opposed by the treatment community and by the New York State Division of Substance Abuse Services—a result that provided an unequivocally negative answer to the question of feasibility. Although in principle it might be argued that addicts are a heterogenous mixture of persons with widely different social needs and therefore that a range of different programs (including special facilities for pharmacological induction and social evaluation during the first month of treatment) are needed for the efficient use of resources, the treatment community and governmental administration have spoken with a virtually unanimous voice: Minimum-service programs will be prohibited, even as preliminaries to full-service programs.

On May 23, 1991, the US Public Health Service working group on methadone treatment rescinded plans for a regulatory change that would have permitted the operation of interim facilities for addicts on waiting lists of licensed clinics. The goal for the future, apparently, is to eliminate the embarassment of waiting lists by discouraging narcotic addicts from applying for treatment. ☐

Vincent P. Dole, MD

Vincent P. Dole, MD, is Professor and Senior Physician, Emeritus, The Rockefeller University, 1230 York Avenue, New York NY 10021-6399. He can be contacted for reprints at the above address.

Reference

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Errata

In: Pass KA, Schedlbauer LM, MacCubbin PA, Glebatis DM. III. Comparison of newborn screening records and birth certificates to estimate bias in newborn HIV serosurveys. *Am J Public Health.* May 1991;81(suppl):22–24.

In table III-1, on page 23, the region labels are reversed. The first two lines of data in the left column should be labeled "NYS Exclusive of NYC"; and the third and fourth lines should be labeled "NYC."

* * *

In: Glebatis DM, Novick LF, Stacy A. VIII. Hospitalization of HIV-seropositive newborns with AIDS-related disease within the first year of life. *Am J Public Health.* May 1991;81(suppl):46–49.

Four lines were omitted at the bottom of column 2 on page 46. The complete paragraph should read: "A total of 995 HIV-seropositive newborns born in December 1987 through June 1988 were confirmed in the Newborn HIV Seroprevalence Study. A total of 151 infants born during this time period were hospitalized with an ARD within 12 months of birth. Assuming that all members of the hospitalized set were members of the seroprevalence set, 15 percent of HIV-seropositive infants born in these seven months were hospitalized with an ARD within 12 months of birth. The monthly ratios ranged from 14 to 18 percent (Table 1)."

Adult Foster Care for the Elderly in Oregon: A Mainstream Alternative to Nursing Homes?

Rosalie A. Kane, DSW, Robert L. Kane, MD, Laurel Hixon Illston, MSPH, John A. Nyman, PhD, Michael D. Finch, PhD

ABSTRACT

Background. In Oregon, adult foster care (AFC) homes, which are private residences where a live-in manager cares for one to five disabled residents, have been covered by Medicaid since 1981 and seem to offer a mainstream alternative to nursing homes. They house almost 6000 older people, two thirds of which pay privately.

Methods. In a cross-sectional study, we interviewed 400 AFC and 400 nursing home residents. Data analyses included descriptive cross-tabulations; hierarchical loglinear models for judging the effects of care setting and payment status on resident characteristics; and logit analyses for predicting care setting and payment status within care settings.

Results. On average, nursing home residents were more physically and cognitively impaired than AFC residents, but there was considerable overlap in patterns of frailty in the two settings. Medicaid AFC residents were less disabled than privately paying AFC residents. AFC residents reported more social activity, even when we controlled for disability status. AFC residents and their families were more likely to value privacy and homelike settings when choosing a care setting, whereas nursing home residents were more likely to value rehabilitation and organized activity programs.

Conclusions. Both AFC and nursing homes are viable components of a long-term care repertoire. The greater disability levels of private-pay AFC residents refutes the criticisms that disabled Medicaid residents were being inappropriately channeled to AFC. (*Am J Public Health.* 1990;81:1113–1120)

Introduction

The repertoire of suggested care to supplant or supplement nursing homes has been growing, although demonstration research thus far has not shown the cost-effectiveness of such care in major randomized trials.[1–3] Lackluster results of demonstration projects are attributed, in part, to poor targeting of the new community care programs to people unlikely to enter nursing homes or to stay there for long. It is also inherently difficult to deliver home care as efficiently as nursing home care to people with heavy care needs. Economies of scale are required, and this, in turn, requires residential proximity of those needing care, which is certainly a characteristic of nursing homes.

This line of reasoning has led to enhanced interest in developing residential settings for functionally impaired persons that permit economies of scale in their care but allow the older disabled persons to preserve personal autonomy and privacy and have more control over their routines. In such living settings, the resident is not viewed first and foremost as a patient.

Study Setting

Adult foster homes, as they have developed in Oregon, provide one such residential setting. An adult foster home in Oregon is a private residence in a residentially zoned area that is licensed to care for one to five disabled adults. Each foster home must have a live-in resident manager (who may also own the home). Meals are served family style, and regulations are kept to a minimum to enhance a homelike atmosphere and autonomy for the residents. The resident manager provides personal care and housekeeping services. If necessary, other services, such as skilled nursing, can be brought into the foster home from home health agencies.

Although the idea of adult foster care is inherently attractive as a concept, several demonstrations projects in other states have failed to develop a strong market for the service.[4–8] However, Oregon's experience with adult foster care has been different in two ways. First, although Medicaid clients were the first to widely use adult foster care, by 1988 private-paying customers constituted more than two thirds of the nearly 6000 residents in foster care. Second, and related, in other states where foster care exists, the programs tend to be proportionately much smaller and largely serve low-income clients with mental illness or developmental disabilities.[9] Only in Oregon does adult foster care seem a mainstream possibility for the elderly.

The growth of adult foster homes in Oregon was stimulated in 1982 when the state made adult foster care a benefit under its Medicaid waiver program. To receive subsidies under this program, the adult foster care resident needed to be eligible for nursing home care on the basis of functional disabilities. The state developed a vigorous program to divert Med-

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Rosalie A. Kane is with the School of Public Health and the School of Social Work of the University of Minnesota, all other authors are with the School of Public Health of the University of Minnesota.

Requests for reprints should be sent to Rosalie A. Kane, DSW, Division of Health Services Research & Policy, School of Public Health, University of Minnesota, 420 Delaware Street, SE, Box 729, Minneapolis, MN 55455.

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icaid clients already in nursing homes to foster homes when the routes to home had been cut off and urged clients who could no longer remain at home to consider adult foster care as an alternative to nursing home care. By 1988, adult foster care was widely accepted by the public as a desirable means of caring for older persons in need of a supportive environment.

Several facts about adult foster care in Oregon were clear. First, the growing utilization suggested that many elderly and families found adult foster care a desirable option. Second, the success of adult foster care within both the private and Medicaid markets (along with vigorous home care programs) was sufficient to reduce the demand for nursing home beds. Between 1981 and 1989, the supply of nursing home beds per 1000 persons aged 75 years and older fell from 50 to 40, and the occupancy rate decreased from 92% to 87%.[10] Oregon is the only state where nursing home use dropped in absolute as well as proportional terms. Third, adult foster care is a less expensive program than nursing home care; both the Medicaid rates and the average private pay rates in foster care were about 2/3 of those in nursing homes. And fourth, adult foster homes became the focal point for criticism from the nursing home industry, which claimed that the care was inadequate, that the degree of social stimulation was insufficient, and that the Medicaid program, in particular, was denying recipients access to the appropriate level of care.

In this context, the state of Oregon commissioned the University of Minnesota to evaluate the adult foster care program. The study had several components, including a mailed census of the foster care industry and a retrospective study of change over time using program data for the Medicaid population. Both of these will be reported in future papers. This paper reports the results of a cross-sectional comparison of 400 foster care residents and 400 nursing home residents. Our analysis looked at differences between care settings and differences between Medicaid and private pay clients within care settings. The ramifications of these questions for public policy extend beyond Oregon at a time when serious consideration is being given to tightening the regulation of group homes and some authorities argue that elderly people should be channeled to particular care settings on the basis of measurable functional characteristics.

Method

Sample

Between December 1988 and February 1989, we conducted in-person, structured interviews with 402 intermediate care facility (ICF) residents and 405 foster home residents in four geographic areas, the latter chosen to represent different facets of long-term care delivery in Oregon (i.e., rural vs urban, Medicaid administration by area agencies on aging vs by regional state offices of the Senior Services Division) while reflecting the average, typical state experience on descriptors such as proportion of elderly, and adult foster home and ICF bed-to-population ratios. Original data collection was necessary because, although considerable data are available about Medicaid residents living in adult foster homes and in nursing homes through their periodic assessments by case managers, little existing data are available on the private pay group in either setting, particularly in foster homes. Moreover, even the Medicaid data set has no information on social activities or choice of care setting.

We generated the foster care sample by randomly selecting foster homes in each geographic area and enrolling all residents of the selected foster homes until we reached the desired proportional number of respondents for each of the four regions. Enough nursing homes were selected randomly in each geographic area to allow us to interview 10 residents in each facility and achieve a number of subjects equal to the foster home sample. From a list of all ICF residents in the randomly selected homes, we randomly selected 5 Medicaid and 5 private pay ICF residents, thus reflecting the proportion of Medicaid and private pay residents in the population. (The nursing home sample was limited to ICF residents because the skilled nursing care level in Oregon is a very small program and widely agreed to include only very disabled and medically unstable persons who are not appropriately compared to adult foster home residents).

To reach the desired sample size of at least 400 residents in each care setting, 108 foster homes were included in the sample; only 4 foster homes approached refused to participate. The final nursing home sample included 39 facilities; only 2 nursing homes approached refused to participate. Participation of the sampled residents was also very high. Actual refusals to participate represented only 4%—6 ICF residents and 5 family members on behalf of a cognitively impaired ICF resident declined to participate. No refusals occurred among foster care residents.

Data Collection

Interviewers were recruited from the local areas and trained by University of Minnesota staff. Most had previous experience in human services or health care. Background information such as date of admission, payment status, medical diagnosis, and family member most knowledgeable about the resident (should a family proxy be needed) was first gathered from the care setting staff on each sample member. For this purpose, records were consulted when necessary. If the resident in either care setting seemed to be an unreliable informant, the interviewers were instructed to skip to a later section of the schedule and attempt to administer the mental status questionnaire. When few of the 10 MSQ questions were answered correctly, or when the resident was clearly too confused to answer, the interviewers moved to proxy respondents. Questions about choice of care setting and satisfaction with care setting were administered by telephone to a family member. The foster care home's resident manager or a nursing home staff member who knew the resident best served as the proxy respondent for questions about current physical functioning, activities, and recent use of hospitals, emergency rooms, and physicians.

Instruments

Identical questions were asked of adult foster home and ICF residents (merely interchanging the terms "adult foster home" and "nursing home" when appropriate.) The functional abilities of the residents were measured through questions carefully designed to be compatible with those found in the SSD 360 (the comprehensive assessment format used by case managers to provide a database on all Medicaid waiver clients in Oregon.) Activities of daily living and instrumental activities of daily living (ADL/IADL) items included toileting, dressing, transferring, bathing, eating, ambulation, and internal medicine use. Cognitive status was measured by the short, portable mental status questionnaire. The health section included self-reported health, vision, hearing, self-diagnosis, and health care utilization. For this study we added three questions to tap mood (depression, anxiety, and positive affect) and questions on activities, recent relationships, satisfaction items, and items on factors related to choice of care setting adapted from an earlier study of nursing home outcomes.[11] The questionnaire and the Oregon SSD 360 are available from the authors. The instrument was characterized by its use of direct questions to respondents rather than by interviewer ratings of the attributes of the respondents.

Statistical Analysis

All data analyzed were categorical. Crosstabs were run on all data with hierarchical loglinear three-way chi-squares used to judge whether the main effects of client setting or payment status or their interactions were significant. Logit analysis was performed to determine which factors predict care setting and payment status within care setting.

Findings

Case-Mix Differences

As shown in Table 1, the demographic characteristics of the samples were similar across care settings, with payment status in foster care reflecting actual proportions in the population. We have separated the private and Medicaid supported clients in all analyses because of policy interest. There were, however, some substantial differences between the two groups. Foster care residents were less likely to come from acute care hospitals. Only 12% and 13% of Medicaid and private pay residents were admitted from a hospital, compared with 44% and 33% of corresponding nursing home patients. Fourteen percent of Medicaid and 17% of private adult foster care residents came from nursing homes, whereas only 9% of Medicaid residents and 7% of private nursing home residents came from adult foster homes.

Table 2 contrasts the pattern of diagnoses. The residents in nursing homes had more major health problems than those in adult foster care; yet this information must be interpreted with some degree of caution because some adult foster care providers had imprecise knowledge of their clients' diagnoses, and the records in adult foster care homes were less complete than in nursing homes. There were also differences by payment status. These did not follow as clear a pattern. For example, whereas clients with dementia were more frequently private pay in adult foster care, they were about equally distributed in nursing homes at a rate still higher. By contrast, patients with seizures were more

TABLE 1—Demographic Background of Sample

Characteristic	No. (%) of Adult Foster Home Residents (n = 405)	No. (%) of Nursing Home Residents (n = 402)
Gender		
Percent Female	314 (78)	293 (74)
Age, y		
Under 50	16 (4)	10 (2)
50–64	21 (5)	12 (3)
65–74	51 (13)	46 (11)
75–84	117 (29)	116 (29)
85–94	156 (38)	164 (41)
95+	45 (11)	54 (13)
Marital Status		
Married	39 (10)	78 (20)
Widowed	260 (67)	249 (64)
Divorced	53 (14)	24 (6)
Single	38 (10)	41 (11)
Education		
Grade school or less	169 (47)	165 (45)
Some or all high school	120 (33)	106 (29)
Some college	50 (14)	57 (16)
College or more	23 (6)	37 (10)
Payment Status		
Private	255 (63)	214 (53)[a]
Medicaid	150 (37)	188 (47)

[a] Although the original intent was to have equal distribution between Medicaid and private pay individuals in nursing homes, two homes drawn in the sample accept only private-pay residents.

TABLE 2—Major Health Problems

	Adult Foster Home		Nursing Home	
Medical Problem	Medicaid, % (n = 150)	Private, % (n = 255)	Medicaid, % (n = 188)	Private, % (n = 214)
Dementia	23	36	42	44
Heart disease	20	22	32	28
Arthritis	13	11	14	10
Stroke	16	23	21	21
Blood pressure	11	5	10	9
Skeletal/muscular	10	10	15	9
Mental illness	14	11	6	7
Vision	8	7	5	4
Hip fracture	3	5	7	17
Respiratory	9	4	10	5
Parkinson's	4	5	5	10
Diabetes	11	8	10	7
Alcoholism/drug abuse	7	1	3	3
Incontinence	4	3	2	3
Mental retardation/developmental disability	5	1	3	—
Old age	3	5	1	—
Hearing	3	5	2	2
Other neurological	3	1	5	5
Cancer	4	2	1	4
Seizures	3	2	9	2
Other bowel and bladder problem	1	1	8	11
Multiple sclerosis	1	1	2	3
Other diagnoses[a]	4	3	12	14
Other nondiagnostic problem[b]	10	7	3	5
Unknown to respondent	4	3	—	—

[a] This category includes any specific diagnosis that was mentioned less than 1% of the time.
[b] This category includes any "nondiagnostic" problem mentioned less than 1% of the time—e.g., weight loss.

TABLE 3—Functional Abilities by Care Setting and Payment Status

	Ability to Function, %											
	Without help				Some Help				Complete Help			
	Foster Home		Nursing Home		Foster Home		Nursing Home		Foster Home		Nursing Home	
ADL Function	Medicaid	Private	Medicaid	Private	Medicaid	Private	Medicaid	Private	Medicaid	Private	Medicaid	Private
Moving	56	48	27	22	34	29	22	25	9	23	52	53
Transferring	86	63	27	24	7	18	20	19	7	19	54	57
Bathing	33	17	6	7	34	32	19	17	33	51	75	76
Dressing	65	48	19	19	22	25	25	20	13	27	56	61
Eating	88	80	60	53	8	13	24	28	4	7	16	20
Toileting	81	65	26	23	12	16	19	17	7	19	54	60

ADL = activities of daily living.

TABLE 4—Cognitive Status by Care Setting and Payment Status

	Adult Foster Home, No. (%)				Nursing Home (ICF), No. (%)			
	Medicaid		Private		Medicaid		Private	
Five or more errors on the SPMSQ	38	(27)	44	(3)	42	(36)	52	(43)
Unable to administer the SPMSQ	11	(7)	58	(23)	67	(36)	87	(41)
Total	49	(34)	122	(56)	109	(72)	139	(84)

SPMSQ = Short Portable Mental Status Questionnaire.

likely to be Medicaid clients in nursing homes and developmentally disabled clients were more likely to be supported by Medicaid in both settings.

The extent of dependency in ADLs is shown in Table 3. More adult foster care clients were able to perform each of the six ADLs studied without assistance. Private paying residents in adult foster care were more disabled than those supported by Medicaid, whereas the proportions were more similar within the nursing home sample. This pattern of difference by payment status is also seen in the levels of cognitive performance on the 10-item Short Portable Mental Status Questionnaire (SPMSQ[12]; see Table 4). Among adult foster care residents, 33% of the private-pay clients had five or more errors, and another 23% could not respond to the test at all. Among the Medicaid foster home group, 27% had five or more errors, and only 7% could not respond. In the nursing home sample, 43% of the private-pay clients made five or more errors, and 41% could not respond; for Medicaid residents the comparable figures were 36% and 36%. Thus a total of 56% of private-pay adult foster care residents and 34% of Medicaid residents did poorly on the SPMSQ compared with 84% of private pay nursing home residents and 72% of Medicaid residents.

TABLE 5—Proportion Considering Selected Factors Important in Selection of Care Setting

	Adult Foster Home, %		Nursing Home, %	
Factor[a]	Medicaid (n = 150)	Private (n = 255)	Medicaid (n = 188)	Private (n = 214)
Homelike atmosphere	94	96	63	62
Safe, supervised place to live	93	96	94	92
Personal assistance	82	84	87	88
Privacy	80	82	54	54
Medical care	80	60	82	80
Flexible routines and rules	81	78	57	51
Cost	70	77	68	66
Location/neighborhood	68	67	69	72
Organized activities	33	26	52	38
Physical rehabilitation program	26	21	51	39
Compatibility or familiarity with facility, staff, or residents	9	4	1	5
Quality care and staff	4	5	8	10
Clean and pleasant surroundings	2	3	8	7
Other	6	10	5	5
None of the above	—	—	2	1

[a]Respondents could select more than one factor as important.

Other reflections of the differences across these settings can be seen in the amount and location of physician care (not shown in a table). Although the proportion of clients reporting a physician visit in the previous 6 months was similar by payment status, nursing home patients were more likely to see a physician. Among adult foster care residents, 80% of Medicaid and 77% of private residents saw a doctor in the previous 6 months compared with 90% of Medicaid and 91% of private-pay nursing home patients. Caution should be used in interpreting these figures because physician contact is mandated on a periodic basis for Medicaid nursing home clients. The location of these visits was quite different. Whereas about 90% of the adult foster care residents saw a physician in his or her office,

TABLE 6—Activities in the Care Setting in the Last Few Months by Functional Status, %

	Never				Sometimes				Often			
	Hi[a]		Lo[b]		Hi		Lo		Hi		Lo	
Activity	AFC	NH	AFC	NH	AFC	NH	AFC	NH	AFC	NH	AFC	NH
Watch television	8	17	35	42	27	39	35	31	65	44	30	28
Listen to radio, to music	43	35	45	28	32	34	30	39	25	32	25	33
Talk to other residents	6	7	39	38	35	47	30	38	59	47	32	24
Read books or newspapers	29	42	69	78	26	24	15	11	45	34	17	11
Engage in hobbies	66	82	94	93	17	12	6	5	17	6	—	2
Chat with staff	3	8	15	25	29	47	35	43	68	45	50	32
Play cards or games	80	57	94	90	14	28	4	8	6	16	2	2
Do activities organized by home	61	28	85	37	26	51	11	44	13	21	4	19
Help with cooking/ housekeeping	68	94	100	98	16	2	—	1	17	3	—	1

AFC = adult foster care; NH = nursing home.

[a] Residents who did not need complete help on more than bathing (n = 290 in foster homes and n = 126 in nursing homes).

[b] Residents who need complete help with four or more functional activities (n = 54 in foster homes and n = 211 in nursing homes).

only about 30% of the nursing home patients did so.

Choice of Care Setting

Mostly, the preferences of those living in adult foster care were similar to those of nursing home residents. As shown in Table 5, however, there were some important differences. For example, adult foster home residents were more concerned about a homelike atmosphere and privacy, whereas the nursing home residents emphasized organized activities and physical rehabilitation. Not surprisingly, both groups valued safety and personal assistance. In choosing one setting over another, 60% of the foster care residents and 36% of the nursing home residents (or their family proxies) believed that they had at least some control over the choice of care setting. A nursing home had been considered by 20% of foster care residents and a foster home had been considered by 15% of nursing home residents.

TABLE 7—Frequency with Which Residents Engaged in Activities Outside the Home, Adjusted for Functional Status in Foster Home and Nursing Home, %

	Hi[a] Status		Low[b] Status	
	AFC	NH	AFC	NH
Frequency				
Not at all	18	48	72	83
Less than once a month	30	26	20	13
Less than once a week	13	11	—	2
About weekly	26	11	7	1
Almost every day	13	5	—	—
Activity				
Take a walk in the neighborhood	36	34	—	11
Use a bus or public transportation	12	17	—	3
Go for a ride in a car	95	92	87	91
Visit relatives or friends	78	80	80	71
Attend church or other religious activities	29	18	—	24
Attend a senior center or club	17	9	7	—
Go shopping	52	34	20	14
Go to a restaurant, movie, or other entertainment	69	59	47	38
Attend activities outside the home arranged by the foster home	31	32	13	23

AFC = adult foster care; NH = nursing home.

[a] Residents who did not need complete help on more than bathing (n = 290 in foster homes and 126 in nursing homes).

[b] Residents who need complete help with four or more functional activities (n = 54 in foster homes and 211 in nursing homes).

Living Conditions

The differences in the two settings can also be seen in the way in which residents spent their time. Table 6 describes the residents' activities according to their level of functioning in their respective care settings. Better functioning adult foster care residents were more likely to watch television, talk with other residents, and help with cooking and housekeeping, whereas better functioning nursing home residents were more likely to participate in organized activities and games and listen to the radio. As shown in Table 7, adult foster care residents were more likely to leave the care setting for activities other than medical appointments, especially shopping and entertainment, a finding that held true when we controlled for functional status and for cognitive status. The reversal of the usual pattern for low functioning nursing home patients reflects the fact that nursing home staff is more likely to organize group activities or to bring activities to the residents. The patterns of responses in Table 8 suggest that adult foster care residents were less likely to get visitors than nursing home residents but more likely to have telephone contact. Nearly 25% of the nursing home residents compared with only about 10% of adult foster care residents had a visitor every day. Only about 5% to 10% of both groups had no visitors. Telephone access was more difficult in nursing homes. About 33% of adult foster care residents com-

TABLE 8—Frequency of Visitors and Phone Contacts in the Last 6 Months by Cognitive and Functional Status in Foster Homes and Nursing Homes, %

	Cognitive Status				Functional Status			
	Hi[a]		Lo[b]		Hi[c]		Lo[d]	
	AFC	NH	AFC	NH	AFC	NH	AFC	NH
Have visitors								
Not at all	6	0	5	9	6	8	6	10
Less than once a month	26	13	24	25	28	16	17	23
Less than once a week	18	17	20	17	18	20	15	16
About weekly	41	32	39	31	39	34	49	29
Almost every day	9	29	12	18	9	22	13	22
Telephone contact								
Not at all	20	50	54	85	26	49	67	83
Less than once a month	22	15	22	7	25	13	9	6
Less than once a week	10	11	9	2	10	12	6	1
About weekly	30	13	10	5	25	15	9	4
Almost every day	17	12	4	2	14	11	9	3

AFC = adult foster care; NH = nursing home.

[a] Residents who answered four or fewer questions on the Short Portable Mental Status Questionnaire (SPMSQ) incorrectly (n = 235 in foster homes and 155 in nursing homes).

[b] Residents who answered five or more questions on the SPMSQ incorrectly or were unable to answer any SPMSQ questions (n = 169 in foster homes and 249 in nursing homes).

[c] Residents who did not need complete help on more than bathing (n = 290 in foster homes and n = 126 in nursing homes).

[d] Residents who needed complete help with four or more functional activities (n = 54 in foster homes and n = 211 in nursing homes).

TABLE 9—Satisfaction among the Cognitively Intact[a] in Foster Homes and Nursing Homes, %

	Very Satisfied		Satisfied		Unsatisfied		Very Unsatisfied	
	AFC	NH	AFC	NH	AFC	NH	AFC	NH
The food	47	23	47	54	4	18	1	5
Your room	46	31	50	61	3	7	1	1
Your roommate	30	23	62	64	8	9	—	4
Your daily care	45	27	52	64	3	9	—	1
Your daily activities	26	21	65	69	8	9	—	1
The amount of time you see family/friends	33	23	49	57	14	15	4	5
Your medical care	39	27	57	63	2	8	3	1
Your physical safety	45	32	53	64	1	3	1	1
The safety of your personal possessions	44	21	53	58	31	16	—	6

AFC = adult foster care; NH = nursing home.

[a] Residents who answered four or fewer questions on the SPMSQ incorrectly (n = 229 in foster homes and 143 in nursing homes).

pared with about 70% of nursing home residents had no telephone contact. Both cognitive and functional status was related to telephone access in both settings, but were less discriminating for visitors.

Satisfaction was generally high in both groups. Table 9 presents data on the reported satisfaction ratings for a number of items among the cognitively intact respondents. In general, the adult foster care respondents expressed stronger levels of satisfaction and less dissatisfaction. The exception to this pattern was in the area of safety of personal possessions where the adult foster care respondents showed greater degrees of dissatisfaction despite a larger proportion also being very satisfied.

Foster care residents and nursing home residents were similar in their use of hospitals in the past 6 months (18% for foster care compared with 15% for nursing home care) though the reported stays for foster care residents were, on average, 1.5 days longer (7.2 compared with 5.7). Foster care residents used emergency rooms more than did nursing home residents (20% compared with 8%).

Predicting Care Setting and Payment Status

We performed discriminant analysis on selected variables to identify which distinguished between adult foster care residents and nursing home residents and to predict payment status within each of the care settings. We have characterized the underlying care needs of our population with selected ADL and IADL variables: mobility, transferring, bathing, dressing, feeding, toileting, a summary score, and taking medicines. It was also hypothesized that a number of diagnoses would play a role in distinguishing between types of residents (residents with heart disease, skeletal-muscular problems, stroke, hip fracture, cancer, diabetes, incontinence, vision problems, hearing problems, and mental illness). To capture cognitive functioning, we included a variable for whether a proxy was used for the interview. We have also included five variables that describe where the client was living or from where he or she was admitted prior to coming to the setting: in a relative's home, in an adult foster home, in a residential care facility, in a nursing home, or from the hospital. Finally, we included two other demographic variables in the analysis: marital status and gender.

Table 10 shows that, of the functioning status variables, transferring, toileting, and taking medicines were significant in explaining the likelihood of a client's entering an adult foster home, whereas none predicted nursing home residency. Women were almost 19% more likely to be in adult foster homes. Residents who had previously lived in their family's home or other foster homes were 15% to 35% more likely to be in adult foster care. Only hip fractures and being on publicly subsidized care explained the likelihood of entering a nursing home.

Two variables, transferring and bathing, were important in estimating that an adult foster care resident is receiving Medicaid (see Table 11). In fact, residents needing help with transferring were more than 38% more likely to be in foster homes than in nursing homes. Marital status alone helped estimate clients with private sources of payment.

Table 12 shows that in a nursing home Medicaid payment status can be predicted by the presence of skeletal-muscular disease and the ability to take medicines. Privately paying residents, on the other hand, are more likely to be married, to have had cancer or a hip fracture as a diagnosis, and be incontinent.

By using hierarchical loglinear models to explain differences between care setting or payment status, we determined that none of our main effects or first level interactions was found to be significant.

Discussion

These data corroborate impressions that, although adult foster care serves many elderly people with disabilities, the two forms of care serve rather different groups of the frail elderly. The patterns of frailty overlap to a considerable degree, but adult foster care treats a population that is, on the average, less physically and cognitively dependent than that found in nursing homes. Part of this effect may be due to the displacement that has already occurred, but at least part of it corresponds to the care needs of two distinct groups, at least at the ends of the distributions. Private-pay residents in foster care, however, show more functional impairment and cognitive impairment than do the Medicaid residents, making implausible the contention that state policy was driving Medicaid residents to accept a lower level of care than they might choose. Because the study is cross-sectional, we cannot say whether this discrepancy is due to more impairment among private-paying clients at admission or to a lesser likelihood that the residents are discharged to a nursing home, or to both.

TABLE 10—Variables That Explain the Likelihood of a Person's Entering an Adult Foster Home[a]

	b	Significance	Percentage[b]
Being female	.79778	<.01	17.1
Having previously lived in relative's home	1.11529	.03	27.1
Having previously lived in adult foster home	1.56264	<.01	37.2
Being Medicaid recipient	−.89297	<.01	−19.9
Being able to transfer	1.18578	<.01	27.3
Being able to toilet	1.32605	<.01	30.5
Being able to take medicines	.674659	.03	16.2
Having hip fracture	−1.22595	<.01	−23.1

[a] The logit analysis was weighted to reflect the actual distribution of adult foster home care in the population, and appropriate standard errors were compiled following Manski and McFadden (1981).[13]

[b] This represents the percent change in the probability of adult foster home placement as opposed to nursing home placement for people having the specified characteristics.

TABLE 11—Variables That Explain the Likelihood of a Person's Being a Medicaid Client in an Adult Foster Home

	b	Significance	Percentage[a]
Having previously lived in nursing home	1.0741	.02	26.1
Being able to transfer	1.6591	<.01	33.6
Being able to bathe	.9386	.01	22.7
Having questionnaire completed by proxy	1.0877	.03	22.3
Being married	−1.1066	.03	−22.0

[a] This represents the percent change in the probability of an adult foster home resident's being supported by Medicaid rather than by private means when these residents have the specified characteristics.

It seems unlikely that adult foster care will replace nursing home care completely, and there will continue to be a need for both services. But adult foster care will probably continue to serve some (and perhaps a growing number) of persons who could also be legitimately served in nursing homes. Categorizing client needs in long-term care remains an imprecise activity, permitting a great deal of discretion in care planning. The policy issue becomes one of whether long-term care will be envisaged as a continuum with a niche for each client or whether it will be viewed as a repertoire of services that allows considerable choice based on lifestyle preferences and prices. A related policy issue cast in relief by this study is whether socially oriented care settings such as adult foster care in Oregon should be more heavily regulated to protect the residents at the risk perhaps of transforming the settings from homes to mini-nursing homes. This situation is exacerbated by the phenomenon of aging in place. Although the admission cohorts to the two forms of care may look more dissimilar, with time they grow increasingly to resemble each other and lead one to conclude that the care is interchangeable. It would, thus, appear both impractical and undesirable to formulate policies based on a rigid delineation of the "appropriateness" of the particular resident for the particular setting. It appears from our data that residents and their families at least in part shop for the setting that is "appropriate."

Moreover, adult foster care has proven attractive to the private market. It is important to appreciate that the growth of this sector in Oregon has not been the result of pressure from the Medicaid agency. Clients and their families able to choose the location of their care have opted for the greater privacy and ambiance and lower price of the adult foster home.

The apparent success of adult foster care has spawned concerns about the need for stricter regulation of this industry on a level more akin to that applied to nursing homes. This pressure is understandable, but it threatens the essence of the program; that is, an approach that relies on an extended family type of arrangement for most care and community agencies for the more technically sophisticated services.

From this cross-sectional study, it was impossible to examine change over time for residents in either care setting, although it appeared that, even when we controlled for disability, foster care residents enjoyed somewhat more social and community activity. In a forthcoming paper, we will compare change in functional outcomes for Medicaid foster care residents and nursing home residents, using the secondary data base available in the Medicaid program. The results of the present analyses offer support to those who are enthusiastic about the adult foster home as a viable option for some subset of those ordinarily served in nursing homes, while recognizing that on average the nursing home residents are more disabled. Indeed, the presence of the foster care program (and other home care programs not discussed in this paper) makes it likely that the nursing home population will be

TABLE 12—Variables That Explain the Likelihood of a Person's Being a Medicaid Client in a Nursing Home

	b	Significance	Percentage*
Having a skeletal muscular problem as diagnosis	.9363	.02	22.9
Having a hip fracture as diagnosis	−1.2999	<.01	−28.2
Having cancer as diagnosis	−2.1125	.02	−36.9
Having incontinence as diagnosis	−.7965	.02	−19.0
Being able to take medicines	1.0284	.05	24.9
Being married	−.6693	.04	−15.9

*This represents the percent change in the probability of a nursing home resident's being supported by Medicaid rather than by private means when the residents have the specified characteristics.

more disabled than nursing home populations in other states. ☐

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International Workshop of Computing in Occupational and Environmental Health to Be Held in Paris in November 1991

The 3rd International Workshop of Computing in Occupational and Environmental Health will be held November 7 & 8, 1991, in Paris, France, at the Hotel Meridien-Porte Maillot.

Professionals in occupational and environmental health as well as computer scientists will compare their latest applications in the following areas:

- individual medical date records and process
- on-line or CD ROM databases development, particularly for bibliography and toxicology
- statistical and epidemiological applications
- management facilitations
- educational uses
- information and knowledge dissemination

As part of the meeting, an exhibition of services and equipment will be provided. For further information, please contact Maryke Lefebvre, SOPREP, 36 rue des Mathurins, 75008 Paris, France. Tel: 331/42.65.17.08.

Is Coronary-Care-Unit Admission Restricted for Elderly Patients? A Multicenter Study

Craig Fleming, MD, Ralph B. D'Agostino, PhD, and Harry P. Selker, MD, MSPH

ABSTRACT

Background. To investigate whether elderly patients are more likely to experience restricted access to high technology medical care, we examined the impact of age on the likelihood of coronary care unit (CCU) admission for patients with acute myocardial infarction.

Methods. As part of a prospective investigation of emergency room triage for patients with suspected cardiac ischemia, we studied 4223 patients presenting to six hospitals. Because CCU admission is the accepted standard of care for acute infarction, we defined nonadmission to the CCU as a restriction of access to care. We used a logistic regression model to control for gender, hospital, and CCU occupancy at the time of admission and examined the relationship between age and CCU nonadmission.

Results. Patients 75 years or older with acute myocardial infarction were 2.5 times more likely not to be admitted to the CCU than younger patients (RR 2.5, 95% CI 1.64, 3.85). Coronary unit admission was restricted even when the physician's admitting diagnosis was "myocardial infarction" (RR 7.1, 95% CI 2.1, 25.0) or "rule-out myocardial infarction" (RR 1.5, 95% CI 1.1, 2.1). Observed differences in clinical presentation or severity of illness between older and younger patients did not account for these findings.

Conclusion. Our findings suggest that physicians may intentionally restrict access to coronary care for elderly patients with acute myocardial infarction. (*Am J Public Health.* 1991;81:1121–1126)

Introduction

As health care expenditures continue to grow, concerns about the cost of medical care for the aged have arisen.[1,2] People over 75 years of age constitute the most rapidly growing segment of our population.[3] Because per capita costs of health care for the elderly exceed those of younger people, further increases in health care expenditures seem inevitable as our population ages. Intensive-care-unit (ICU) care, which accounts for a substantial portion of hospital-related outlays, is considered a prime target for limiting health care expenditures.[4]

In response to these concerns, society and the medical profession have begun to question the appropriateness of unrestricted use of high technology medical care in the elderly.[5–7] While there is presently no consensus regarding appropriate limitations of such care, it has been suggested that the threshold to admit patients to ICUs may already be age-related, resulting in less frequent admission of older patients despite more severe disease among the elderly.[8] To examine this issue, we studied the impact of age on physicians' coronary care unit (CCU) admitting practices for emergency room (ER) patients with symptoms suggesting acute cardiac ischemia. Our objective was to determine if differences existed in CCU admission practices for older patients, as compared with younger patients, that could not be attributed to differences in clinical presentation or severity of illness.

Methods

The data on which this study was based were collected prospectively during a multicenter clinical trial to study CCU admitting practices and to develop and test an instrument predicting the likelihood of acute ischemic heart disease among patients in the ER setting. More detailed information regarding the data collection and design of the study may be found in the paper describing that study.[9]

Participating Hospitals

The six participating hospitals in the study included two urban medical-school teaching centers, two medical school-affiliated teaching hospitals in smaller cities, and two rural nonteaching hospitals.

Study Subjects

The study included all consenting patients, 40 years of age and older, presenting to the ERs of any of the six hospitals, with a chief complaint of chest, jaw, or left arm pain, shortness of breath, or other symptoms suggestive of acute ischemic heart disease. These criteria were based on the Imminent Myocardial Infarction Rotterdam Study and identify more than 90 percent of patients with acute cardiac ischemia who seek medical care.[10] Patients were asked to participate in a study of the diagnosis and treatment of heart disease, and among patients meeting the entry criteria, 92 percent consented to par-

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From the Center for Cardiovascular Health Services Research and Study Design, New England Medical Center Hospitals and Tufts University School of Medicine (all authors); and the Department of Mathematics, Boston University (D'Agostino); Dr. Fleming is currently at Dartmouth Medical School.

Address reprint requests to Harry P. Selker, MD, MSPH, Center for Cardiovascular Health Services Research Unit, New England Medical Center, 750 Washington Street, Boston, MA 02111. This paper, submitted to the Journal April 12, 1990, was revised and accepted for publication January 23, 1991.

TABLE 1—Physicians' Triage Disposition for Younger and Older Emergency Room Patients with Acute Myocardial Infarction, Angina Pectoris, and without Acute Ischemic Heart Disease (N = 4,223)

	Triage Disposition						
	Admitted to Coronary Care Unit		Admitted to Ward		Discharged Home		Total
Final Diagnosis	n	%	n	%	n	%	n
Myocardial Infarction							
Age 40–74	557	89.8	45	7.3	18	2.9	620
Age ≥ 75	146	76.8*	42	22.1	2	1.1	190
Angina Pectoris							
Age 40–74	377	66.6	137	24.2	52	9.2	566
Age ≥ 75	82	57.0	51	35.4	11	7.6	144
No Acute Ischemic Disease							
Age 40–74	566	27.8	605	29.7	868	42.5	2,039
Age ≥ 75	183	27.5	373	56.2	108	16.3	664

*Relative risk (RR) of not being admitted to the CCU for older patients compared with younger patients is 2.50 (95% CI = 1.64, 3.85)

TABLE 2—Physicians' Triage Disposition for Patients with Acute Myocardial Infarction by Age Groups.

Age (years)	Patients Admitted to CCU n (%)	Patients Not Admitted to CCU n (%)	Relative Risk CCU Nonadmission (95% CI)
40–74	557 (90)	63 (10)	1.0
75–79	69 (79)	18 (21)	2.09 (1.16,3.37)
80–84	50 (81)	12 (19)	2.18 (1.09,4.36)
85+	27 (66)	14 (34)	4.25 (2.10,8.59)

Relative risks are adjusted for admitting hospital and compare rates of CCU nonadmission for older age groups to those of patients in the 40–74 year age group.

ticipate in the study. Patients were enrolled during a 16-month period ending in August 1981. For the present study, we excluded all data collected during the experimental phase of our previous clinical trial of CCU use.[9]

Data Collection

Clinical data, cardiac-enzyme tests, and electrocardiograms (ECGs) were obtained at the time of entry for all consenting patients. Additional clinical data—including clinical records, follow-up ECGs, and cardiac-enzyme tests—were obtained for both hospitalized and nonhospitalized patients at 48 hours after admission. Mortality was determined from study inclusion through follow-up, including the index hospitalization for hospitalized patients. The overall follow-up rate was 89 percent.

Data Analysis

Study patients were stratified by age into a younger group (40–74 years) and an older group (75 years of age or older). Final diagnoses were assigned for each patient based on blinded review of the serial ECGs, results of the cardiac-enzyme tests, and the ER and hospital discharge records. Final "true" diagnoses of acute myocardial infarction or angina pectoris were based on clinical presentation, ECGs, and cardiac-enzyme tests using the World Health Organization[11] and New York Heart Association criteria.[12] We assigned final diagnoses on the basis of the patient's ER clinical presentation. Patients admitted to the hospital with other diagnoses, who subsequently had myocardial infarctions in hospital, were analyzed according to their initial diagnosis. For patients with acute myocardial infarction, clinical severity was stratified by Killip class at presentation.[13]

The probability of not being admitted to the CCU (i.e. CCU nonadmission) was the primary end-point of this study. For the purposes of this study, CCU care and ICU care are considered equivalent. "Step-down" Units were not included as CCU care because such units represent a less intensive standard of care. At the time of this study, CCU care was the accepted standard of care in the United States for patients with acute ischemic heart disease.[14] Hence, CCU nonadmission is considered to be an adverse event for patients with acute myocardial infarction.

Statistical Methods

All hypothesis tests were two-tailed. The hypothesis that physicians' CCU admitting practices were related to patient age (Table 1) was tested using a logistic regression model predicting the probability that the patient was not admitted to the CCU. Separate models were analyzed for each group of final verified diagnoses (acute myocardial infarction, angina pectoris and no acute ischemic heart disease). The factors included in each model were: age, gender, the hospital where the patient was evaluated (to control for differences in admitting rates among hospitals), and the occupancy level of the CCU at the time of the patient's admission. Patients over 75 years of age with acute myocardial infarction were also stratified by five-year age intervals (Table 2) and a similar logistic regression model was used to calculate age-specific relative risks for CCU nonadmission.[15]

The test for linear trends in proportions[16] was used to examine trends in CCU nonadmission rates for patients with myocardial infarction when stratified by Killip Class (Table 3). To test the hypothesis that differences in physicians' CCU admitting practices for patients with acute myocardial infarction were related to differences in clinical presentation between older and younger patients (Table 4), we used a multivariate logistic regression model predicting CCU nonadmission. The factors included in the model were: age, gender, the hospital where the patient was evaluated, presenting features, and separate interaction terms to detect relationships between age and each of the presenting features. The interaction terms were used to examine whether advanced age and the presence of a particular clinical sign or symptom had a significant impact on CCU nonadmission.

Tests for the observed differences in CCU nonadmission rates for patients on the basis of the ER physician's admitting diagnosis (Table 5) were performed using the Mantel-Haenszel Chi-square procedure to control for differences in hospital admitting rates.[17]

To examine the relationship between mortality and CCU nonadmission for older versus younger patients with acute myocardial infarction, we used a logistic regression model predicting mortality. The factors included in the model were: age, gender, the hospital where the patient was evaluated, CCU nonadmission and an interaction term for the relationship between age and CCU nonadmission. The interaction term was used to examine whether CCU nonadmission had a significant impact on the mortality of elderly patients with acute infarctions.

Results

The average sample size for the six hospitals was 704 patients and ranged from 453 to 938 patients. By hospital type, 1,505 patients were seen in the urban medical school teaching hospitals, 1,309 in the medical school-affiliated teaching hospitals in smaller cities, and 1,409 in the rural non-teaching hospitals.

The study included 3,225 patients (76 percent) ages 40 to 74 years, and 998 patients (24 percent) who were 75 years of age and older. The mean age was 58.9 years among the younger group of patients, and 81.0 years among the older patients. Male patients constituted 58 percent of the younger patients versus 41 percent of the older patients. Of all study patients presenting to the ER, 810 patients (19 percent) had acute myocardial infarction, 710 patients (17 percent) had angina pectoris, and 2,703 (64 percent) proved to have neither form of acute ischemia. No significant differences were found between younger and older study patients in the proportion of patients with acute myocardial infarction (19 percent in each group) or angina pectoris (18 percent vs 14 percent, respectively).

Although not presented here, the triage dispositions of patients 65–74 years of age did not significantly differ from those of younger patients, reinforcing the inclusion of this group with the middle-aged group. Age 75 years was selected as best representing a truly elderly group of patients, as noted in other discussions of health care for the elderly.[1,5]

Among ER patients who proved to have acute myocardial infarction (Table 1), CCU nonadmission occurred 2.50 times more frequently for patients 75 years of age or older than for patients 40 to 74 years of age (95% CI = 1.64, 3.85). Examination of the differences in CCU nonadmission rates among ER patients with acute myocardial infarction, stratified by five-year age intervals for older patients, shows that the likelihood of CCU nonadmission increases with advancing age (Table 2).

Among ER patients with angina pectoris (Table 1), the probability of CCU nonadmission for older patients compared to younger patients was less evident (RR 1.45, 95 percent CI = 0.98, 2.13). For patients without acute ischemic heart disease, CCU nonadmission rates for older and younger ER patients were not significantly different, although older ER patients were significantly less likely to be discharged to home from the ER than younger patients (RR 0.26, 95 percent CI = 0.21, 0.32).

Differences in the clinical severity of acute myocardial infarctions by Killip classes at the time of presentation for older and younger patients are shown in the first set of columns in Table 3. A clear trend towards an increasing proportion of more serious infarctions in older patients is seen. The second set of columns in Table 3 shows the differences in CCU nonadmission rates for older versus younger patients in each presenting Killip classes. For all presenting Killip classes, except Killip class 4, CCU nonadmission was more likely among older patients than younger patients. Among study patients with Killip class 4 myocardial infarctions, defined as presentation with cardiogenic shock, only one patient was not admitted to the CCU. The apparent trend towards increasing CCU nonadmission rates with advancing Killip class for older patients with Killip class 1, 2, and 3 myocardial infarctions is also clear.

To test the hypothesis that older patients with infarction were admitted to the CCU less frequently than younger patients due to less clearly diagnostic presenting signs and symptoms, we examined gender, ECG findings, and presenting complaints of chest pain, dyspnea, or other symptoms (Table 4). Two ECG findings indicative of acute ischemic cardiac disease, T-Wave peaking or inversion and ST-segment elevation or depression greater than 1 mm, increased the likelihood of CCU admission. Aside from age, only presentation with dyspnea and no chest pain predicted a decreased likelihood of CCU admission. We found no significant interactions in the regression equation between advanced age and any other presenting feature. Thus, older patients with any specific presenting feature were neither more nor less likely to be admitted to the CCU than younger patients with similar presentations. Furthermore, controlling for the effects of all of

TABLE 3—Distribution of Acute Myocardial Infarctions and CCU Nonadmission by Presenting Killip Classes among Younger and Older Emergency Room Patients (N = 810)

	Patients with Acute Myocardial Infarction n (%)		CCU Nonadmission for Patients in Each Killip Class n (%)	
Killip Class	Age 40–74 (n = 620)	Age ≥ 75* (n = 190)	Age 40–74 (n = 63)	Age ≥ 75** (n = 44)
1	429 (69)	80 (42)	48 (11)	14 (18)
2	107 (17)	47 (25)	9 (8)	11 (23)
3	62 (10)	48 (25)	5 (8)	19 (40)
4	22 (4)	15 (8)	1 (5)	0 (0)

*p < 0.0001 for trend of increased proportion of older patients with higher Killip classes.

**p < 0.0001 for trend of decreased CCU admission of older patients with higher Killip classes.

Percent of patients not admitted to the CCU are calculated relative to the numbers of patients in each age group within that Killip class.

TABLE 4—Multivariate Logistic Model of Presenting Features Associated with Nonadmission to the Coronary Care Unit Admission of Emergency Room Patients with Acute Myocardial Infarction (N = 810)

Presenting Feature (n)	Relative Risk for Nonadmission to the CCU RR (95% CI)
Age ≥ 75 (190)	2.4 (1.5, 3.8)
Male (503)	0.9 (0.5, 1.4)
Chest Pain in the ER (693)	1.6 (0.7, 4.0)
Both Dyspnea and Chest Pain (199)	1.1 (0.6, 1.9)
Dyspnea and No Chest Pain (78)	4.5 (1.8, 11.1)
No Dyspnea and No Chest Pain (56)	0.4 (0.1, 1.3)
ECG: T-Waves Peaked or Inverted (540)	0.5 (0.3, 0.7)
ECG: ST Elevation or Depression (681)	0.4 (0.2, 0.6)

No statistically significant interactions were found between age and other presenting features. NS—feature is not a significant predictor of difference in triage disposition at the p = 0.05 level.

TABLE 5—Triage Disposition for Younger and Older Emergency Room Patients by Emergency Room Physician's Admitting Diagnosis (N = 4,223)

	Emergency Room Patients with Each Diagnosis n		Patients with Each Diagnosis Not Admitted to the CCU n (%)	
Admitting Diagnosis	Age 40–74 (n = 3,225)	Age ≥ 75 (n = 998)	Age 40–74	Age ≥75
"Myocardial Infarction"	225	59	5 (2)	8 (14)
"Rule-out Myocardial Infarction"	1,116	294	223 (20)	79 (27)
All Other Diagnoses	1,854	645	1,502 (81)	497 (77)

Percentages of patients not admitted to the CCU are calculated relative to the numbers of patients in each age group with the diagnosis.

the presenting features on CCU nonadmission in a multivariate logistic model had no significant impact on the risk of age as an independent negative predictor of CCU admission.

To provide further support for the hypothesis that differences in triage disposition of elderly patients did not result simply from misdiagnosis, we examined the admitting diagnoses of the ER physicians to find whether differences in CCU admissions occurred for older and younger patients even when the ER physician believed the patient to have acute ischemic heart disease (Table 5). Among all ER patients with an admitting diagnosis of "myocardial infarction," 86 percent of patients 75 years or older were admitted to the CCU, compared with 98 percent of younger patients. Expressed as a relative risk, these older patients were 7.1 times more likely not to receive CCU care (95% CI = 2.1, 25.0). For ER patients with an admitting diagnosis of "rule-out myocardial infarction," older patients were 1.5 times more likely not to receive CCU care (95% CI = 1.1, 2.1). Further, upon review of the available 28 ER medical records for the 44 older patients with acute infarction not admitted to the CCU, there was no documentation that the patients themselves had requested a less intense level of care, or that they had impaired mental or social function.

Finally, did differences in triage disposition for older and younger patients have a measurable impact on mortality? Among all patients with diagnosed acute infarctions, mortality was greater for older patients than younger patients (29 vs 12 percent), but was similar among older patients admitted to the CCU versus those admitted to the wards or sent home (RR 1.12, 95% CI = 0.38, 3.7).

Discussion

The results of this study suggest that advanced age has an impact on the CCU admitting decisions made by physicians, most evident among ER patients with acute myocardial infarctions. None of the hospitals participating in this study had an explicit policy of rationing intensive care to elderly patients and lower CCU admission rates for older patients with acute infarction was not related to the occupancy level of the CCU. The observed difference in CCU admission rates for older patients with acute infarction did not result simply from diagnostic misclassification in the ER. Nevertheless, our data do not allow us to directly observe the reasoning or motivation underlying the triage decision.

Physicians' awareness that a study of CCU admitting practices was taking place may have affected our results even though physicians were not aware of the study hypothesis. Differences in the triage disposition of older and younger patients might have occurred if physicians improved their diagnostic accuracy more in younger patients than in older patients. However, this fails to explain why physicians, even when they had made a diagnosis of acute ischemic heart disease, as indicated by their admitting diagnosis of "myocardial infarction" or "rule-out myocardial infarction," were still less likely to admit an older patient to the CCU.

This study was primarily designed to examine differences in CCU admitting practices, and despite its size, its power to detect small differences in mortality and other outcomes among different subgroups of patients is limited. This is reflected by the fairly wide confidence intervals for the relative risk of mortality in older patients based on CCU admission. Because Killip classes are correlated with mortality,[13] the observation that among patients not admitted to the CCU, older patients had higher Killip classes of infarctions than younger patients suggests that differences in mortality may indeed exist.

Several factors not included in our data set may influence the ER physicians' decision not to admit elderly patients to the CCU. First, physicians may believe that CCU care for elderly patients is not cost-effective, particularly since unequivocal evidence demonstrating the efficacy of CCUs is lacking[18,19] and there is some evidence that medical intensive care is probably overutilized in the United States.[19] However, the differences in triage disposition we describe occur among patients with acute myocardial infarction, for whom CCU care is the present standard of care regardless of age. Thus, the notion that the physicians in our study were somehow utilizing CCU care more effectively for older than younger patients does not provide an appropriate explanation for our findings.

A second concern is that CCU care may be potentially harmful for older patients, because of the association between age and an increased likelihood of iatrogenic complications.[20] Thus, physicians may intentionally restrict CCU care for older patients for compassionate reasons.

Third, mentally competent patients, by exercising their acknowledged rights of self-determination, may themselves influence treatment decisions by refusing more intense levels of care.[21] However, review of available ER medical records for older patients with acute infarction in our study, who were not admitted to the CCU, yielded no evidence that patients themselves had influenced the admitting decision.

Fourth, socioeconomic factors may affect the level of intensity of care given to patients. For example, marital and insurance status have recently been shown to influence the choice of therapy for patients with lung cancer.[22]

Finally, it has been shown that the patient's capacity to interact meaningfully with others in established social roles influences physicians' decision making in the management of critically ill patients.[4] Thus, physicians might intentionally restrict CCU admission for elderly patients with dementia or severe disability. In our study, all participating patients were able to give informed consent and provide necessary information to the research assistants, so it is unlikely that dementia among older patients could account for the observed differences in CCU admission. Moreover, review of the available charts of elderly patients not admitted to the CCU, despite having admitting diagnoses of "myocardial infarction" or "rule-out myocardial infarction," yielded no evidence that such considerations played a role in this decision.

Several previous studies have examined the influence of age on ICU admission practices. In a study conducted at an urban hospital, it was noted that, although 10 percent of the general population age 75 years and older reside in nursing homes, only 5 percent of the ICU/CCU patients were nursing home residents.[23] In another study, the age distribution of patients receiving respiratory intensive care was found to be similar to the age distribution of the general hospital population, except for patients over 70 years of age, among whom there were fewer ICU admissions.[24] In a third study, examining selected patients admitted to the ICU and the general medical wards at the same hospital, no difference was found between the age distributions of ward patients and ICU patients, although there was a borderline significant higher severity of illness for older medical ward patients.[25] Overall, these studies, which looked at patterns of care at single institutions, provide indirect evidence of selective triage to ICU care for older patients.

This prospective multicenter study, following patients from their presentation in the ER, provides more direct evidence of restrictive use of intensive care for the elderly. If age influences the likelihood of ICU admission, then studying ICU patients alone, or even ICU along with ward patients, will introduce inclusion bias in study subject selection. To properly interpret data concerning the benefit or access to ICU care for elderly patients, it is important to study the "denominator" group of all patients seen in the emergency room who were candidates for ICU or ward admission, rather than "numerator" patients in one or two specific sites within the hospital.

The fact that the data for this study were collected between 1979 and 1981 might be considered a limitation to the usefulness of our findings. However, an advantage of these data is that they were collected prior to the advent of Medicare's prospective payment system. Provision of intensive care, a costly part of hospital care, is susceptible to PPS incentives,[4] which might confound a study of physician or CCU admission practices done more recently.

The data on which our findings are based also precede the widespread use of thrombolytic therapy for patients with acute myocardial infarctions. Most randomized trials of thrombolysis have excluded patients over the age of 70 because of fear of increased complications in the elderly.[26] If the availability of thrombolytic therapy creates an incentive to preferentially admit to the CCU those patients who receive thrombolysis, this may create another bias against CCU admission in elderly patients, particularly when limited CCU resources are available.[19] As hospital payment policy evolves and ongoing and planned trials of thrombolytic therapy that include aged patients are completed, these influences may be mitigated.

We have presented data suggesting that advanced age has an impact on the triage disposition of patients presenting with suspected acute ischemic heart disease that is not explained by differences in severity of illness or varying clinical presentations among older patients. While our findings are not definitive proof, they suggest that physicians may intentionally restrict access to coronary care for elderly patients. These findings should prompt further investigations to find consistent or countervailing evidence for elderly patients in other intensive care settings and with other conditions. ☐

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Migrant Farm Workers in Need of Better Health Care

The Migrant Clinicians Network recently released the results of *Migrant Health Status: Profile of a Population With Complex Health Problems*, the first comprehensive investigation into health conditions for migrant and seasonal farm workers. Specific to those farm workers who migrate within the midwestern migratory stream, the study reveals an overall health status with higher rates of infectious disease than in the general population.

The results from this study are shocking. The demographic patterns, socioeconomic conditions, life-style characteristics, and disease categories reflect an agrarian society in a Third World country caught in an infectious disease cycle. Factors such as poverty, malnutrition, infectious and parasitic diseases, poor education, a young population, and poor housing equate to a highly vulnerable population in need of resources. Clearly, the migrant population is at greater risk and suffers more problems than the general population. The results of this research demonstrate the need for more service, care, and treatment. The need for developing a health policy and research agenda for migrant farm workers in this decade is evident.

The Migrant Clinicians Network (MCN) is a multidisciplinary organization of all health and human service professionals who work with migrant and seasonal farmworkers. The network was established in 1984 to identify and address clinical issues affecting the health status of the nation's migrant farmworker population. MCN comprises migrant clinicians from 41 states and Puerto Rico, representing a wide range of professional disciplines and areas of expertise. The Network serves as a clinical forum for migrant and seasonal farmworker health issues.

The health status study was conducted under the direction of G. E. Alan Dever, PhD, of Mercer University. For a copy of the executive summary, study monograph, or complete report contact the Migrant Clinicians Network, 2512 South IH35, Suite 220, Austin, TX 78704. Tel: 512/447-0770.

Comparison of Subjective Ratings of Function with Observed Functional Ability of Frail Older Persons

Janet T. Elam, BS, Marshall J. Graney, PhD, Terinell Beaver, MD, Douaa El Derwi, MD, William B. Applegate, MD, MPH and Stephen T. Miller, MD

ABSTRACT

Background. Important clinical decisions often hinge on patients' functional status. Previous studies have shown disagreement among sources of ratings of patients' functional status. This study compared patient self-ratings, family member ratings, and physician ratings of patient function to performance-based functional testing criteria.

Methods. Five activities of daily living of 73 older patients were studied at admission to a rehabilitation unit following discharge from an acute care community hospital. Data were collected from patients, family members, and physicians and were compared with performance-based function testing.

Results. Patient ratings were significantly more accurate than physician ratings for walking, transferring, and telephoning. Patients were significantly more accurate than family members for rating walking and telephoning, but patients were not significantly more accurate than family members or physicians for rating eating or dressing.

Conclusions. We conclude that decisions about patients' functional level should be based on performance testing. If performance testing is unavailable, patients' own ratings are most accurate, followed by family ratings. Physicians' ratings are least accurate. (*Am J Public Health.* 1991;81:1127–1130)

Introduction

Functional assessment of elderly patients has been a focus of research for more than 25 years.[1] Its importance is demonstrated by the one fifth of persons aged 65 and older in the United States who are unable to perform at least one measure of activities of daily living (ADL).[2] The complexity of functional measurement is demonstrated by more than 40 published ADL scales.[3] Most ADL scales use patient self-report to rate performance of tasks. Recently investigators have compared ADLs measured by an assessment instrument or geriatric evaluation to patient self-report, family report,[4] or physician report[5] and found that different sources of ADL rating may not agree and therefore are not interchangeable.

A recent review by Guralnik and colleagues[6] makes a strong case for the value of performance-based ADL measures. Elderly patients who experience an acute illness or injury often have decreased function requiring decisions ranging from assistance with personal care to permanent nursing home placement. During hospitalization patients may be unable to assess their own functional status because their activities may be medically limited. Thus, accurate assessments of function are needed to ensure appropriate recommendations. Because sources of ADL ratings may disagree, comparisons of different rating sources with objective measures of performance are important to characterize the accuracy of various rating sources.

This study was designed to evaluate performance-based ADL testing in a community hospital rehabilitation unit. Unlike previous studies, we compared the accuracy of self-ratings, family ratings, and physician ratings to performance-based measures.

Methods

Study Sample

Seventy-three patients aged 60 years and older were studied between January 1987 and March 1989. These patients had been hospitalized at the Baptist Memorial Hospital, a large community hospital, for acute medical or surgical problems and required additional rehabilitation.[7] Patients were eligible for enrollment in the study if they agreed to participate, signed informed consent, were cognitively able to participate (no more than four errors on the Short Portable Mental Status Questionnaire),[8] were not aphasic, and had a family member currently involved in their care. This was defined as a relative who visited the patient during the acute care hospital stay and was willing to assume responsibilities for assistance when the patient was discharged.

Measurement

Although the number of days of the acute hospital stay was not recorded, all

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Janet T. Elam, Douaa El Derwi, William B. Applegate, and Stephen T. Miller are with the Department of Preventive Medicine, and Marshall J. Graney is with the Departments of Preventive Medicine and Biostatistics, University of Tennessee Memphis. Terinell Beaver, formerly with the University of Tennessee, is now in private practice.

Requests for reprints should be sent to Janet T. Elam, Department of Preventive Medicine, University of Tennessee Memphis, Memphis, TN 38163.

This paper was submitted to the journal May 25, 1990, and accepted with revisions March 15, 1991.

TABLE 1—Demographic and Functional Health Data (N = 73)

Variable	Data
Age (years)	
Mean ± SD	77.3 ± 7.9
Range	61–98
Sex (% female)	62.5
Race (% white)	87.7
Marital status (% married)	31.5
Baseline ADLs (% independent)[a]	
Walking	42.5
Transferring	52.8
Telephoning	77.5
Eating	93.1
Dressing	77.5

[a] Observed performance test data.

patients enrolled in the study were medically stable and ready for rehabilitation. As part of the initial evaluation at transfer to geriatric rehabilitation (usually the same day) patients were first asked to rate their abilities to walk, dress, eat, transfer, and use the telephone, using a standard format from the Framingham Disability Study.[9] Five performance levels were possible: (1) independence, (2) requires use of equipment, (3) requires assistance of another person, (4) requires assistance of another person and equipment, or (5) unable to perform. These same abilities were rated separately by the caregiving family member and by the patient's personal physician. At admission patients were also asked to perform five tests which corresponded to the five rated abilities: (1) walking; walk 25 feet; (2) dressing: button a shirt; (3) eating: transfer dried beans with a spoon from one dish to another; (4) transferring: transfer from bed to chair; and (5) telephoning: make a telephone call.

Interviews and performance-based testing were conducted by a nurse clinician or by one of two physician research fellows. Family members were interviewed at the rehabilitation unit or by telephone. All interviews with patients' personal physicians were made by telephone. Interviews with family members and physicians were made within 48 hours of the patient interview or considered unobtainable. Demographic information was collected from patient's medical records.

Patient interviews and observational protocols were completed for 73 patients. Sixty-four (88%) family member and 49 (67%) physician interviews were completed within 48 hours of the time patient data were collected. Complete observational, patient, family, and physician data collection protocols were executed for 41 (56%) patients. Data on observed ADL, age, sex, race, and marital status were analyzed to detect potential bias caused by some physicians' nonparticipation. No statistically significant differences in these variables were found between patients whose physicians participated and those whose physicians did not.

Data Analysis

Data were analyzed using the SPSS^{x} system of computer programs, including CONDESCRIPTIVE for descriptive statistics, CROSSTAB for truth tables, and PEARSON CORR for correlations.[10] The χ-square test was used to determine the independence of ratings vs the performance test criterion. Differences between these χ-square tests were used to test for differences in the agreements of ratings with the criterion. In all cases, $P<.05$ was used as the criterion of statistical significance. Significance levels were adjusted for multiple comparisons when appropriate through use of the Bonferroni method.[11]

For analysis of accuracy of ADL ratings the five levels of ability for each ADL were dichotomized as independent vs dependent. Independence was defined as performance of the ADL without assistance from another person, and dependence was defined as nonperformance or need for the assistance of another person to perform the ADL.

Actual performance ability, assessed by direct observational testing, served as the criterion of performance of an ADL. Criteria for evaluating diagnostic tests (including sensitivity, specificity, predictive value of a positive test, and predictive value of a negative test) were adopted to evaluate the accuracy of ratings by use of the analogy whereby functional dependence was equated with the presence of a disease condition, i.e., a positive test. Thus, sensitivity was defined as the number of patients rated dependent who also were observed to be dependent divided by the number observed to be dependent. Specificity was defined as the number of patients rated not dependent who also were observed to be not dependent divided by the number observed to be not dependent. Predictive value of rated positive was defined as the number rated and observed to be dependent divided by the number rated dependent. Predictive value of rated negative was defined as the number rated and observed to be not dependent divided by the number rated not dependent.

Results

Demographic and functional health data for the 73 study patients are presented in Table 1. Patients were transferred to rehabilitation with diagnoses representing all major rehabilitation groups. Twenty-four percent had a primary diagnosis of stroke, 17% had experienced fractures (of these 50% were hip fractures), 48% had a primary medical diagnosis, and 11% had other problems.

To investigate overall rater performance in accurately documenting patient ADL status, zero-order correlations of patient, family, and physician ratings for each ADL compared to observed performance are presented as phi coefficients in Table 2. For walking, the ratings of patient, family, and physician were all significantly correlated to observed performance. For transferring, telephoning, and dressing, both patient and family ratings were significantly correlated to observed performance. For eating, only patient ratings were significantly related to observed performance. Thus, patient ratings were significantly correlated to performance data for five of five ADL, family ratings for four of five, and physician ratings for one of five.

Other results comparing observed performance to patient, family, and physician ratings are also shown in Table 2. For walking and transferring, physician ratings have low specificity (walking=45%, transferring=38%) due to underestimation of the number of independent patients. Patient self-ratings ($P<.005$) and family ratings ($P<.005$) were both significantly more highly correlated to observed performance than were physician ratings for walking and transferring. In addition, patient self-ratings for walking were significantly more highly correlated to observed performance than were family ratings ($P<.005$).

For telephone use, physician ratings had low sensitivity (25%) because physicians rated too few patients as unable to use the telephone. Family raters also had difficulty with rating telephone use, and their underestimation of dependency was documented by the family raters' low predictive value of rated positive (42%). Patient self-ratings ($P<.005$) and family ratings ($P<.05$) were both significantly more highly correlated to observed performance than physician ratings for telephone use. As with walking, patient self-ratings of telephone use were also significantly more highly correlated to observed performance than family ratings were ($P<.005$).

For eating, patient, family, and physician too often rated patients as dependent, resulting in poor predictive values of rated positive. In addition, physician ratings lacked sensitivity, not detecting which patients were dependent in eating. All raters lacked specificity and showed poor predictive values of rated positive for dressing, chiefly because of underrating independent dressers.

Discussion

When an elderly patient has an acute illness or accident, decisions about the level of care required at hospital discharge can permanently alter the patient's living arrangement. Questions arise concerning need for postdischarge services such as home health care. Although on the one hand it is advisable to avoid inappropriate institutional placements, on the other hand patients living in the community who are too functionally dependent to care for themselves pose a burden for themselves and others.

How best to evaluate functional level is a question with which researchers have long struggled. Self-reported ADLs may include potential that the patient has not yet regained or represent performance that the patient may be unable to attain at the time of observation because of temporary physical status or environmental limitations. Our experience with performance testing supports the opinions of Guralnik et al.[6] that performance testing can be easily administered, is not overly time consuming, is well received by patients, and results in no harm to patients. Although we did not use the same tests, our methods for measuring function were similar to those described by Guralnik and colleagues. Also, the general agreement of patient and family ratings with observed function data validates performance testing.

Our data showed that patients and, to a lesser degree, family members had good knowledge of the patient's ability to walk, transfer, use the telephone, and dress. Also, patient ratings of eating ability were fairly accurate. By type of function, the predictive value of rated positive was highest for patient and family for the lower extremity skills, i.e., walking and transferring. For upper extremity skills the predictive value of rated positive was uniformly lower.

TABLE 2—Estimates of Predictive Accuracy

ADL and Rater	Sensitivity, %	Specificity, %	Predictive Value of Rated: Positive, %	Predictive Value of Rated: Negative, %	Correct, %	Phi Coefficients[a]
Walking						
Patient	90.5	90.3	92.7	87.5	90.4	0.80*
Family	83.3	67.9	76.9	76.0	76.6	0.52*
Physician	82.8	45.0	68.6	64.3	67.3	0.30**
Transferring						
Patient	76.5	81.6	78.8	79.5	79.2	0.58*
Family	90.3	67.6	71.8	88.5	78.5	0.59*
Physician	73.9	38.5	51.5	62.5	55.1	0.13
Telephoning						
Patient	56.2	92.7	69.2	87.9	84.5	0.53*
Family	61.5	78.4	42.1	88.9	75.0	0.35***
Physician	25.0	87.5	33.3	82.4	75.0	0.14
Eating						
Patient	80.0	71.6	17.4	98.0	72.2	0.28**
Family	75.0	71.4	14.3	97.8	71.6	0.23
Physician	40.0	78.6	18.2	91.7	74.5	0.14
Dressing						
Patient	100.0	41.0	23.4	100.0	50.0	0.31****
Family	100.0	35.3	25.0	100.0	46.8	0.30**
Physician	87.5	35.3	21.9	92.3	42.2	0.17

[a] Statistical tests of association between rater and performance data.
*$P < .001$.
**$P < .05$.
***$P < .005$.
****$P < .01$.

Physicians underestimated independence for lower extremity ADLs and overestimated independence of ADLs requiring cognitive skills, such as dressing and telephoning. These findings are in agreement with earlier findings of Pinholt et al.,[5] who found that mental status deficits were underreported in a comparison of ratings of residents, interns, and nurses to functional ratings by geriatricians using assessment instruments. Also, our finding that patient and family ratings were significantly different from each other for walking and telephone use agrees with findings published by Rubenstein et al.,[4] who observed significant ADL report differences between patient and family.[4] In the Rubenstein study, patients and family ratings were compared to ratings by nurses who provided daily patient care. This criterion differs from the criterion in our study, which was performance-based ADL testing. As in Rubinstein's study, our family raters underestimated independence in walking and transferring; however, in contrast to Ruberstein's results, our patient raters had the best overall accuracy. Our study differs from previous studies comparing clinical ADL rating to other rating sources in a fundamental way. Our physician raters were all community hospital based private practitioners, responsible for the patient's daily care during the acute care hospitalization. Most were internists who had a long-term relationship with the patient. This is in contrast to the Rubenstein study, which was set in a Department of Veterans Affairs Hospital,[4] and to the Pinholt study, which was set in a military hospital.[5]

This study represents an attempt to learn about the measurable components of ADLs. Choosing the optimum surrogate measure to use in performance testing is difficult, and some of the differences observed between performance testing and subjective ratings may be due to the fact that the performance measures do not assess all the complex aspects of a given ADL. Variables like walking and transferring can be observed directly. For other ADLs such as dressing and grooming, which require multiple functional and cognitive skills, choosing surrogates is more difficult. For example, our choice of shirt buttoning as a surrogate for dressing may not have been optimal because buttoning lacks representation of lower extremity dressing and does not represent the cognitive processes necessary to select clothing and to don garments in order.

This study has some limitations. Study patients were not representative of the entire population of acute care patients because all had family members involved in their care and all were eligible for rehabilitation. Demented patients may not report their function accurately and were excluded from our study. Our study did not observe continence and toileting, and these ADLs were not included in our ratings. Personal physicians were asked to answer immediately to our inquiry and were not given an opportunity to collect additional information from observations, patients, or family. However, they were the attending physicians during the hospitalization and had ample opportunity to collect these data. All study physicians practiced in a large community hospital in a metropolitan area. These results may not apply to other service settings and may not, for example, represent the situation in nonmetropolitan areas, where physicians might have a lifetime relationship with the patient.

In summary, physicians would be well advised to formulate recommendations that are dependent on function after obtaining objective measures of performance. If performance testing is unavailable, patient self-rating usually serves well for ADLs, especially those related to ambulation. For patients incapable of giving reliable ratings, family rating is the best substitute. If physicians would include questions concerning performance in these areas in their routine history-taking, their knowledge of their patients' performance might be improved. ☐

Acknowledgments

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Call for Abstracts for APHA Late-Breaker Epidemiology Exchange Session

The Epidemiology Section will sponsor a Late-Breaker Epidemiology Exchange on Wednesday, November 13, 1991, at APHA's Annual Meeting in Atlanta. The exchange will provide a forum for presentation of investigations, studies, methods, which have been conceived, conducted, and/or concluded recently so that abstracts could not meet the deadline for submission to other epidemiology sessions. Papers submitted should report on work conducted during the last 6 to 12 months.

Abstracts should be limited to 200 words; no special form is required. Abstracts should be submitted to Robert A. Gunn, MD, Division of Field Epidemiology, Epidemiology Program Office, Mailstop C08, Centers for Disease Control, Atlanta, Goergia 30333 (Tel: 404/639-3661; FAX: 404/639-3950, and must be received by October 4, 1991). Receipt of abstract will be acknowledged by postcard and acceptance will be decided by October 9, 1991 (please include a return envelope).

Screening Recommendations for the Elderly

Mark H. Beers, MD, Arlene Fink, PhD, and John C. Beck, MD

ABSTRACT

Background. Studies have documented the potential contributions of preventive health care programs. Yet little is known about which screening tests should be included in public health programs for older persons. This study offers recommendations regarding these tests.

Methods. The recommendations come from synthesizing the findings of the US Preventive Services Task Force, the literature, and the consensus of experts in geriatrics, gerontology, and health policy research. The literature was evaluated to identify methodologically sound studies of the prevalence of selected disorders and benefits and availability of screening procedures for those disorders. Experts from various fields specializing in the care of the elderly formed panels to assist in evaluating the literature and providing further information from gerontological and public health perspectives.

Results. We recommend vision testing for refractive error; inspection of the skin surface for fungal infection and skin cancer, drug eruptions, and xerosis; a history for symptoms of xerosis; audiometric testing for presbycusis; surveys for hearing loss; otoscopic inspection for cerumen impaction; dental examination for caries; measurement of blood pressure for hypertension; and breast examination and mammography for cancer.

Conclusions. Our study suggests that these screening procedures are useful for public health screening programs. More information is needed on the effects of screening services on the health and functioning of older persons. (*Am J Public Health.* 1991;81:1131–1140)

Introduction

The American nation is growing older.[1] During the past two decades, the population over 65 has grown more than twice as quickly as the rest of the US population. In 1986, 41% of the elderly were 75 years of age and older; by the year 2000, half the elderly are expected to be over age 75. The 85-and-over population is growing especially rapidly. This group is expected to be seven times as large by the middle of the next century as it was in 1980. The number of functionally impaired elderly is already large and is expected to grow rapidly.[2] As the population ages and becomes more impaired, concern for the prevention of disease and promotion of good health increases.[3]

Studies have documented the potential contributions of preventive health care programs.[4] Advocates assert that preventive programs would result in a healthier and more industrious older population that could assist in reducing the nation's expenses for medical care and social support and contribute to the gross national product.[5] Investigators have documented a high prevalence of undetected, correctable medical conditions among older adults living in the community.[6] As many as 94% of elderly persons screened in a community program were found to have some positive finding requiring advice or information.[7]

Little information is available, however, regarding the screening tests that should be an integral component of a public health program for older persons. Unlike screening in office practice, public health programs do not focus on treatment but, instead, emphasize referrals and prevention. Many, like the California Preventive Health for the Aged Program, are publicly funded, use trained nurses, and are mobile, thus limiting their use of some screening equipment. Among the considerations for screening services for public health programs would be the accuracy of such tests in detecting prevalent, asymptomatic disease, the availability and appropriateness of use of screening, and the availability of beneficial treatment of the diseases that are discovered.

We synthesized findings from the literature, the US Preventive Services Task Force,[8,9] and experts in geriatrics, gerontology, and health policy research to identify screening services that are appropriate for public health programs conducted by nurses.

Methods

Data Sources

Expert Panels. We identified laboratory screening tests, periodic medical review, hypertension, mental health, dermatological problems, ophthalmological problems, dental problems, and hearing problems as areas of greatest concern. Using these as a starting point, we consulted with experts in geriatric medicine and gerontology to help us select specific topics (e.g., stool guaiacs as a laboratory screening test or glaucoma as an ophthalmological condition).

We assembled 24 members of UCLA's academic faculty and clinical ex-

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The authors are with the School of Medicine and the School of Public Health, University of California, Los Angeles.

Requests for reprints should be sent to Dr. Mark H. Beers, UCLA, Center for Health Sciences, School of Medicine, Multicampus Division of Geriatric Medicine and Gerontology, Los Angeles, CA 90024.

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perts into eight panels corresponding to the areas listed above. Panel members were selected for their knowledge of preventive health care and of the elderly. Panels ranged in size from four to six members, with at least one physician-member of UCLA's Multicampus Division of Geriatric Medicine and Gerontology and one health services researcher with expertise in geriatrics and consensus methods. Other members were expert in one or more of the study topics.

We adopted a two-round consensus method that has been used successfully to resolve issues of uncertainty in health care.[10,11] During the first round, panelists worked independently and without discussion to complete a written questionnaire. They sent the results for synthesis and then met to discuss the findings. Areas of agreement and disagreement were discussed, with the major emphasis on resolution of disagreements. After discussion, the panelists rerated each question.

Literature Review. Our review included studies that were published in English between 1978 and January 1989. We relied on a selective list of investigations considered by experts to be important in guiding geriatric health care practice rather than striving for a review of all published studies.

Inclusion criteria. The purpose of the literature review was to synthesize current knowledge from methodologically sound studies regarding the incidence and prevalence of selected undetected disorders in the elderly and the effectiveness of screening and early treatment. By screening, we meant any methods of detecting disease, regardless of stage or severity; prevention meant primary and secondary measures. Tertiary prevention was not considered in this project. The methods of screening included physical examination and history; patient questionnaires; instruments such as the sigmoidoscope and ophthalmoscope; and laboratory tests such as fasting glucose and thyroid function tests.

Exclusion criteria. We did not review essays or methodologic studies (e.g., to determine the psychometric properties of a measure of function or mental status), although we examined their reference lists for data sources appropriate to our review. We did not include articles whose main purposes were to investigate the treatments and outcomes of care unrelated to screening. We excluded review articles, except in the case of hypertension, because the results of research had been extensively synthesized.

Sources of information. We identified literature through several mechanisms. All members of the expert panels recommended research studies on screening, and references in those papers were also reviewed. We conducted a MEDLINE search simultaneously.

Abstraction of information. We reviewed all articles suggested by the project members, located through secondary sources, or uncovered by the computerized search (see Appendix A), but we abstracted only those that met our inclusion criteria (see Appendix B). We developed an abstraction form for recording the contents of each article. Ten articles were reviewed by three members of the team to establish the reliability of the abstraction process. We focused on the extent to which the inclusion and exclusion criteria discriminated among articles and whether agreement existed on their contents. When we were satisfied with the form's capability to facilitate consistent results, we each reviewed articles; every fifth article received a second review to ensure the continued uniformity of the abstraction process.

Evaluation of the literature. The literature was evaluated on the basis of the quality of the research design using methods similar to those of the Canadian Task Force on the Periodic Health Examination,[12] the US Preventive Services Task Force,[8] and the RAND/UCLA Variations Study.[13] The best evaluations were assigned to studies with the most rigorous designs (e.g., randomized, controlled trials). Only studies that provided convincing evidence through methodological rigor (e.g., controlled studies) were deemed acceptable.

US Preventive Services Task Force. The results of the US Preventive Services Task Force became available in 1989. The credentials of the Task Force panel members and the methods used to arrive at recommendations are described elsewhere.[8] Our study could not rely solely on the results of the Task Force for three reasons. First, the panelists did not explore all the topics in which we were interested, such as functional assessment. Second, although some Task Force recommendations directly affect older persons, the elderly are not the primary focus, and no geriatricians served on the Task Force's 20-member panel. There are many topics, such as the evaluation of dementia or the prevention of functional impairment, for which the perspective of geriatricians and gerontologists may differ from that of other clinicians and researchers. The third reason for not relying solely on the Task Force is that we were interested in public health programs rather than case management by clinicians. The Task Force addressed some diseases that occur almost exlusively in the elderly, such as glaucoma, and it also made specific recommendations for the elderly regarding hearing impairment and bacteriuria. This information was used as a major source of data in this study.

The Synthesis of Data

Our final recommendation for each topic is based on all data sources from which information was available. We assumed that a screening test should receive the highest recommendation if its use is supported by the literature and the US Preventive Services Task Force, it is not currently used appropriately, and it can effectively uncover prevalent, undetected disease for which beneficial treatment is available.

Throughout this project, we used the risk-benefit definition of appropriateness, meaning that a screen is appropriate when its potential benefit outweighs its potential risks. We based our assessment on the effectiveness of screening tests, i.e., the benefit for a given medical problem under average conditions of use.[14]

In this project, available meant that no barriers (such as lack of coverage by Medicare or scarcity of personnel or equipment) existed to prevent patients from receiving a test or treatment. We also included attention to compliance in our definition of availability because patients might not avail themselves of tests or treatments (e.g., hearing aids) if a social stigma or extreme discomfort were associated with them. By prevalent, we meant that the disease or condition occurred in a high enough percentage of elderly persons to warrant screening. We defined undetected disease as a disease or condition that is unknown to the patient or his or her doctor. A test or treatment was determined to be appropriate if those people who should be receiving it were receiving it. We called a treatment beneficial if it could help a substantial proportion of persons with the disease or condition while harming relatively few.

To arrive at our recommendations, we scored each topic (e.g., dental caries, screening for diabetes) on seven items: (1) extent of support from the US Preventive Services Task Force; (2) extent of support from the consultant experts; (3) extent of support from the literature; (4) prevalence of asymptomatic disease associated with the topic; (5) benefit of treatment for the disease; (6) current availability of treatment; and (7) current use of screening. We used a 9-point scale, with 1 = strongly negative (i.e., low prevalence, and little or no support, benefit, or availability), 5 = equivocal (i.e., average, neutral, or uncertain support, prevalence, or availability), and 9 = strongly positive (i.e., high prevalence, and much support, benefit, and availability). The final score for each topic was the mathematical average of ratings on all seven items. The strength of our recommendations is based on this final score according to the following scale. 9 = strongly recommended, 8 = recommended, 7 = recommended with reservation, 6 = equivocal, 5 = equivocal, 4 = equivocal, 3 = not recommended with reservation, 2 = not recommended, 1 = strongly not recommended. For example, a particular screening test might be strongly recommended by the US Task Force (9), and strongly recommended by the UCLA experts (9), although there is little or no literature supporting its use (5). The disease might be common, occurring in over 10% of the population (8). Moderately effective treatment (7) might exist, but the availability of that treatment might be limited due to its high price that is not reimbursed under Medicare and the reluctance of many individuals to undergo the treatment (2). Without new programs to encourage use of this screening procedure, it might be unavailable currently to most elderly persons (8). Such a scenario would receive a final score of 6.9, which we would round off to 7. We support the use of tests that score 7 or better; tests scoring between 4 and 6 are equivocal; and tests scoring between 1 and 3 are not recommended.

Results

Data Sources

Expert Panels. Topics discussed by the panels and their opinions follow.

Hearing

Hearing loss is extremely common in late life, affecting 25% to 35% of older persons, and it increases dramatically with advancing age. Mild hearing loss is more likely to be undetected. The most common causes of hearing loss in late life are presbycusis, cerumen impaction, and noise-induced sensory neuronal hearing loss. The best screening test for presbycusis and noise-induced hearing loss is pure-tone audiometry; the best screening test for cerumen impaction is otoscopic visualization of the auditory canal. Beneficial treatment exists for all three causes of hearing loss, although amplification is sometimes less successful for noise-induced hearing loss than for presbycusis. The use of amplification is limited because of its cost and the social stigma associated with using hearing aids.

Laboratory screening tests

Approximately 2% to 3% of elderly persons have undetected disease that can be diagnosed by routine laboratory tests. The most common undetected diseases (and the best screening methods for detecting each) are anemia (hematocrit or hemoglobin), bacteriuria (urinalysis), hyperlipidemia (serum cholesterol), diabetes (fasting blood sugar), bleeding colonic lesions (stool guaiac), hypothyroidism (TSH), and electrolyte disturbances caused by diuretics and nonsteroidal anti-inflammatory agents (serum electrolytes).

Beneficial treatments exist for hypothyroidism, bleeding colonic lesions (including early cancer), diabetes, and iatrogenic electrolyte disturbances. The benefits of treating anemia depend on the cause. The benefits of treating hypercholesterolemia probably depend on the age of the patient: treating elevated cholesterol in the very old may not be appropriate. Treating asymptomatic bacteriuria is probably beneficial in men but not in women. There is some limitation to the availability of treatment of diabetes due to the cost of medications and monitoring equipment and compliance with diet, medications, and monitoring. Treatment is limited for iatrogenic complications of medication use because physicians are often reluctant to change medical therapy on the advice of nonphysicians or because medical conditions necessitate their continued use. Hyperlipidemia treatment is limited due to the cost and side-effects of medication and compliance with dietary restrictions and exercise programs.

Periodic medical review

Over 50% of the elderly have some undetected disease or condition. The percentage that could be detected through medical history and limited physical examination could not be determined; however, the most common undetected diseases and conditions (and the best screening method for detecting each) that are likely to be found through periodic review are colon cancer (stool guaiacs), breast cancer (breast examination and mammography), coronary artery disease (history), urinary incontinence (history), inappropriate use of medication (medication review), impaired social functioning (screening scales), impaired physical functioning (screening scales), falls (history), and nutritional problems (weight and height). In almost all cases, these tests and procedures are either not currently available to the majority of elderly persons or are not used appropriately at the current time.

Beneficial treatment exists for early colon cancer, early breast cancer, and coronary artery disease. Treatments for the other diseases and conditions are limited at best, and, in some cases, attempts at treatment may be harmful. Treatment availability is limited for urinary incontinence, correcting inappropriate use of medication, and improving poor social and physical functioning, nutritional problems, and falls.

Ophthalmological disorders

Approximately 10% of elderly persons have undetected eye disease or visual impairment, and the prevalence increases with age. The most common diseases and conditions are glaucoma, cataracts, macular degeneration, and uncorrected refractive error; the oldest-old (persons over 85) are most likely to suffer from macular degeneration.

The panel suggested that when screening is performed by nonophthalmologists or nurses, tonometry is the best screening for glaucoma because it is inexpensive, administration takes minimal training, and expensive or bulky equipment is not required. Testing visual acuity is the best screening procedure for all of the other diseases mentioned. Persons with impaired visual acuity should be referred to an eye specialist for futher evaluation. The minimal barriers to obtaining these procedures are mostly due to lack of reimbursement by some Medicaid programs. Visual acuity testing is generally used appropriately, but tonometry is generally not.

The treatment for glaucoma is generally available and beneficial; treatment for cataracts is available and beneficial; treatment for macular degeneration is not beneficial and, if laser treatment is performed inappropriately or incorrectly, may be harmful; treatment for uncorrected refractive error is beneficial and generally available, limited only by cost and reimbursement.

Mental health

Depression can be defined by strict criteria as stated in the DSM-III-R[15] or by a looser definition of affective disorder serious enough to impact on functioning and well-being. Using the strict definition, 1% to 3% of elderly persons suffer from unrecognized depression, but using the looser definition, 7% to 11% are affected. Few data are available on either the best methods for screening for depression in the elderly or the accuracy of the screening. The panel suggested that the eight-item Medical Outcome Study scale[16] might be useful, especially if followed by the Diagnostic Interview Schedule[17], screening version. The CES-D[18] or Zung[19] scales might be useful alternatives. None of these screening tests is available or used appropriately at this time. Treatment of depression is beneficial, although its availability is limited by toxic side-effects of medication and lack of resources for counseling and other psychological therapies.

Nearly 4% of the elderly suffer from unrecognized dementia or chronic cognitive impairment, although we recognize that in some populations the prevalence may be higher.[20] The prevalence of these conditions increases with age. The panel suggested that the Folstein Mini-Mental State[21] is the best screening procedure, although this is only useful in English-speaking populations. Although these tests are available, they are not currently used appropriately.

Currently no good treatment is available for 80% of elderly persons suffering from cognitive decline; the remaining 20% may be helped by treating reversible illness, including depression. The availability and benefit of treatment are highly dependent on the exact cause and could not be estimated overall.

Dermatological disorders

More than half of elderly persons have undetected dermatological disorders severe enough to lead to complications or complaints, with a higher incidence in sunny states such as California. The percentage of undetected disorders that do not result in the seeking of care could not be estimated. The most common diseases and conditions in this category are fungal infections, skin cancer, drug eruptions, photoaging, and xerosis. Visual inspection by a trained observer is the best screening procedure, although a history of symptoms of dry skin is also important for finding cases of xerosis.

TABLE 1—Characteristics of the Research Studies (N = 33)

Characteristic	No. of Documents	Reference Number[a]
Topics		
Hearing disorders	4	7, 6, 31, 32
Presbycusis		
Cerumen impaction		
Noise-induced loss		
Laboratory screening tests	4	1, 14, 15, 24
Diabetes		
Breast cancer		
Bacteriuria		
Hypercholesterolemia		
Anemia		
Colon cancer and bleeding colonic lesions		
Hypothyroidism		
Electrolyte disturbances		
Periodic medical review	0	—
Coronary artery disease		
Colon cancer		
Breast cancer		
Urinary incontinence		
Inappropriate use of medication		
Impaired social functioning		
Impaired physical functioning		
Nutritional problems		
Falls		
Hypertension	[b]	—
Ophthalmological disorders	0	—
Visual acuity		
Glaucoma		
Macular degeneration		
Cataracts		
Mental health	8	2, 8, 10, 11, 23, 28, 29, 33
Depression		
Dementia and cognitive impairment		
Dermatological disorders	4	3, 26, 27, 30
Skin cancer		
Xerosis		
Photoaging		
Fungal infections		
Drug eruptions		
Dental disorders	13	4, 5, 9, 12, 13, 16–22, 25
Xerostomia		
Caries		
Oral cancer		
Gingivitis and periodontitis		
Issues		
Effective screens	11	1, 3, 8, 10, 11, 14, 17, 23, 24, 30, 32
Subject characteristics	12	3–6, 11, 12, 19, 21, 25, 26, 28, 29
Rates of use	3	9, 18, 22
Incidence/prevalence	21	1, 3–5, 7, 13, 16–21, 25–27, 30
Supporting Agency		
Federal	16	3, 4, 9, 10, 12–14, 16, 18, 19–21, 29
Foundation	2	6, 22
University	3	26, 30, 31
Unidentified	6	7, 17, 23, 25, 27, 33
Starting date of data collection		
Before 1978	2	2, 13
1978–1979	4	1, 15, 23, 24
1980–1983	9	17–21, 25, 29
1984–1989	3	26, 27, 32
No exact dates given	15	3, 6–12, 14, 16, 22, 28, 30, 31, 33
Duration of study		
Under 1 year	1	23
Up to 1 year	6	1, 4, 18, 19, 21, 27
One to 2 years	3	15, 26, 32
2 or more years	6	2, 10, 14, 24, 25, 28

(Continued)

TABLE 1—Continued

Characteristic	No. of Documents	Reference Number[a]
Geographic location of sample		
National with Calif	2	2, 8
Regional without Calif	3	3, 11, 28
Calif statewide	1	27
One state (not Calif)	3	9, 17, 29
Local	16	4–7, 12, 15–22, 26, 30–33
Other	2	9, 17
Sampling		
Random	7	4, 5, 19–21, 27, 28
Population	2	1, 18
Data collection methods		
Record review	8	2, 7, 12, 14, 21, 24, 25, 27
Patient interview/ questionnaire	14	3, 5, 6, 8–11, 15, 17, 18, 22, 28, 29, 33
Provider survey	2	12, 33
Physical exam	14	1, 4, 5, 13, 16, 17, 20, 21, 23, 26, 30–32
Analysis by age		
	17	2–5, 12, 16–22, 25, 28–31
Study site		
Community	24	1, 3–6, 11–13, 15–22, 25–30, 32, 33
Chronic care institution	5	7, 14, 23, 24, 31
Age		
Includes elderly	2	1, 11
Elderly emphasized	11	1, 3, 7, 10, 12, 13, 17, 25, 26, 28–30
Exclusively elderly	17	4, 5, 6, 9, 14–16, 18–22, 24, 27, 31–33
Population/Ethnic composition		
White	23	2–8, 10–12, 16–23, 27, 29–31, 33
Hispanic	6	2, 3, 8, 10, 11, 27
Black	9	2, 3, 6, 8, 10, 11, 22, 27, 29
Asian	4	2, 8, 22, 27
Population/Income		
Middle	2	12, 15
Lower	3	9, 22, 24
Mixed	12	2, 6–8, 10, 14, 17, 28–32
Not stated	16	1, 3–5, 11, 13, 16, 18–21, 23, 25–27, 33

[a] Reference numbers refer to bibliography listing in Appendix B, not the the general reference list.
[b] Only review articles were included. See text for references.

Treatments for fungal infections, skin cancer, drug eruptions, and xerosis are both available and beneficial. The treatment of photoaging is probably not beneficial and generally not available, due to the cost of the medications and side-effects such as excessive skin dryness and pruritus.

Dental disorders

Nearly 50% of the elderly have undetected dental disease. The prevalence of undetected disease is likely to be highest for those with impaired mobility or limited finances. The most common diseases and conditions are caries, advanced gingivitis and periodontitis, xerostomia, and oral lesions including oral cancer. Clinical examination is the best screening procedure in all cases.

Beneficial treatments exist for caries, and, with certain limitations, for periodontitis, xerostomia, and oral lesions. The benefit of treatment of oral lesions depends on its characteristics. Treatment is available for each condition, although the availability of treatment for xerostomia is somewhat limited.

Literature Review. We reviewed 139 studies, 33 of which met our criteria for abstraction. We reviewed five pertaining to ophthalmology and three on periodic review of medical care but did not abstract data from them because they failed to meet our criteria. For most topics, the literature contributed limited data or no data at all.

Table 1 is a complete list of topics studied and the articles abstracted. Our review of the published literature disclosed a paucity of data on the effectiveness of screening. About two thirds of studies focused on the incidence and prevalence of disease, and the remainder emphasized the effectiveness of screening tests. Half the studies were supported by federal sources and one quarter used data collected between 1980 and 1983. In the 16 (48%) studies that specified duration, ten (33%) collected data for less than 2 years.

Most studies relied on local samples rather than statewide or national groups. Seven assessed randomly chosen samples, but none of these was a randomized controlled trial. Approximately 40% used physical examinations as the main source of data. More than half analyzed findings in terms of the age of the sample, and about half focused exclusively on the elderly. In 22 (66%) studies, participants were White.

Ten important review articles were identified that had examined the subject of screening for hypertension in the elderly.[22–31] Hypertension has been established as a risk factor for many serious problems, such as stroke, coronary artery disease, congestive heart failure, and multi-infarct dementia. Although some controversy still remains as to whether treating isolated systolic hypertension decreases mortality and morbidity, the general consensus is that treating both systolic and diastolic hypertension are useful and that detecting asymptomatic hypertension in the elderly is advantageous.

US Preventive Services Task Force. The Task Force provided data on 13 of this study's 34 specific topics. In some cases, the Task Force addressed issues that affect the elderly, but it did not specifically address whether screening needs were different in elderly populations. For four topics (hearing impairment, bacteriuria, elevated cholesterol, and visual acuity), the Task Force did specifically address the elderly, and in another three topics (glaucoma, dementia, and oral cancer), it reviewed topics that affect the elderly almost exclusively, although it did not specifically comment on the elderly. The Task Force recommended screening for hearing loss, elevated cholesterol, and loss of visual acuity.

Recommendations

Table 2 describes this study's recommendations and the bases on which they were made. We recommend the following without reservation: vision testing for refractive error; inspection of the skin surface for fungal infection, skin cancer, drug eruptions, and xerosis; interview for history of xerosis symptoms; audiometric testing for presbycusis and survey scales for hearing loss; otoscopic inspection for cerumen impaction; dental examination for caries; measurement of blood pressure

TABLE 2—Recommendations

Topic	Screen	PR	LT	PN	RXB	RXA	SC	TF	Recommendation[a]
Hearing loss									
Presbycusis	Audiometry	+	++	++	++	<>	<>	+	7
	Self-assessment scale	+	<>	+	++	<>	+	+	7
Cerumen impaction	Otoscopic inspection	++	<>	++	++	+	<>	<>	7
Noise-induced	Audiometry	<>	+	+	+	<>	<>	+	6
Screening laboratory tests									
Diabetes	Fasting blood sugar	+	−	+	+	++	<>	<>	6
Breast cancer	(see Periodic medical review)								
Bacteriuria	Urine analysis	++	−	+	<>	++	+	<>	6
Elevated cholesterol	Serum cholesterol	<>	−	+	<>	<>	−	+	5
Anemia	Hemoglobin or hematocrit	<>	−	+	<>	++	−	−	5
Bleeding colonic lesions	(see Periodic medical review)								
Hypothyroidism	T4	<>	<>	+	++	++	<>	<>	6
Iatrogenic electrolyte disturbances	Serum electrolytes	<>	−	+	+	<>	<>	<>	5
Periodic medical review									
Coronary artery disease	History	<>	<>	+	+	+	<>	<>	6
Colon cancer	Guaiacs, sigmoidoscopy	<>	+	+	+	+	<>	<>	6
		<>	+	<>	+	+	<>	<>	6
Breast cancer	Breast exam, mammogram	+	+	+	+	+	<>	++	8
	—	+	+	+	+	++	<>	+	8
Urinary incontinence	History	+	<>	+	<>	<>	<>	<>	5
Inappropriate use of medications	Medication review	+	<>	+	++	<>	<>	<>	6
Social dysfunction	Assessment scales	+	<>	+	<>	−	+	<>	6
Physical dysfunction	Assessment scales	<>	<>	+	<>	<>	<>	<>	6
Nutrition	Height and weight	<>	<>	+	+	−	<>	<>	5
Falls	History (including safety)	<>	<>	+	<>	<>	<>	<>	5
Hypertension	Sphygmomanometer	+	++	++	+	+	+	+	8
Ophthalmological disorders									
Glaucoma	Tonometry, perimetry, ophthalmoscopy	+	<>	+	+	+	<>	−	6
		+	+	<>	+	+	<>	−	6
		+	<>	−	+	+	<>	−	5
Refractive error	Vision testing	++	<>	++	++	+	−	+	7
Cataracts	Vision testing	+	<>	+	+	++	−	<>	6
Macular degeneration	Vision testing	<>	<>	−	−	<>	−	<>	4
Diabetic retinopathy	Ophthalmoscopy, history and	<>	<>	−	+	+	<>	<>	5
	referral	<>	<>	++	+	+	−	<>	6
Mental health									
Depression	Screening questions	+	<>	+	<>	<>	<>	<>	6
Dementia	Mental status questions	<>	<>	+	<>	<>	<>	<>	6
Dermatological disorders									
Skin cancer	Visual inspection	+	+	+	+	++	<>	+	8
Xerosis	Visual inspection and history	+	<>	+	++	++	<>	<>	7
Photo aging	Visual inspection	+	<>	+	<>	<>	<>	<>	6
Fungal infection	Visual inspection	<>	<>	+	++	++	<>	<>	7
Drug eruptions	Visual inspection	<>	<>	+	++	++	<>	<>	7
Dental disorders									
Xerostomia	Clinical exam	<>	<>	++	+	<>	−	<>	6
Caries	Clinical exam	++	++	++	++	<>	−	<>	7
Oral cancer	Clinical exam	<>	<>	++	+	<>	−	+	6
Gingivitis and periodontitis	Clinical exam	+	<>	++	+	<>	−	<>	6

Note: PR = disease prevalence; LT = literature review; PN = recommendations of panels; RXB = treatment benefit; RXA = treatment availability; SC = current availability of screening; TF = US Preventive Services Task Force. − = negative support; <> = equivocal support or no data; + = supportive data; ++ = strongly supportive data.

[a] 9 = Strongly recommended; 8 = Recommended; 7 = Recommended with reservation; 6 = Equivocal; 5 = Equivocal; 4 = Equivocal; 3 = Not recommended with reservation, 2 = Not recommended; 1 = Strongly not recommended. The prevalence of disease was estimated by panelists when data were not available. The benefits and availability of treatment and the current availability of treatment were rated by panelists.

for hypertension; and breast examination and mammography for cancer.

Discussion

The recommendations presented here are applicable to public health screening programs that aim to refer large numbers of community-dwelling elders and rely on nurses who often travel to deliver services. This study focused on screening for programs staffed by public health nurses, so some tests receiving low ratings might be appropriate for programs relying on physicians. The ophthalmoscope, for example, is an effective screening device for glaucoma when used by a trained ophthalmologist.

From the perspective of clinical geriatrics, many of the tests and procedures that are not supported by this study may still be appropriate when performed by physicians and other health professionals and used for purposes other than screening. Recording mental status, for example, may provide useful baseline data for determining changes in mental status. Also, asking about symptoms of depression, which are often missed by primary care practitioners,[32] is generallly considered important in primary care geriatric medicine. Measuring renal function with serum creatine may also result in useful baseline data, and following serum electrolytes can be useful in persons on medications, such as nonsteroidal anti-inflammatory drugs, when renal function is known to be impaired. Reviewing medication use is almost certainly important in geriatric care when performed by a primary care physician, although screening for inappropriate medication use by others cannot be supported by our research. Thus, recommendations made here for the exclusion of certain measures from screening programs should not be interpreted to mean that these measures should be excluded from good clinical geriatric care.

We limited our investigation of screening to eight broadly defined topics, and these do not encompass all the needs of the elderly. We did not explore screening for vaccination status, for example, or for review of gastrointestinal problems or joint disease. A complete screening package might require other procedures or questions not discussed here. Moreover, the information on which we based our recommendations may change at any time. New research or alterations in reimbursement policies for hearing aids or dental work, for example, would necessitate changes in our ratings. If changes in reimbursement policies were to make hearing aids available to more persons, our ratings for audiometry would move from recommended with reservation to recommended or strongly recommended.

Better data are needed to guide screening programs, and research is needed on such topics as the response of patients and physicians to referrals from screening programs and their outcomes. We need, for example, to determine whether depression screening scales can be used efficiently and accurately in a population of elderly persons, whether physicians are responsive to referrals about depression, and whether patients receive beneficial treatments sooner because of screening. Research is also needed on ways to get patients to use their medications correctly. If inappropriate use (including noncompliance with therapy, use of interacting drugs, or use of drugs that might impair function or cognition) is uncovered, what should be done to alter patients' and physicians' behavior? In fact, additional research on screening is needed for each of the topics explored in this study.

We urge cautious confidence in applying our recommendations. Despite their limitations, they represent the best possible efforts of some leading experts in geriatric health care to weigh scientific and experimental evidence. It is obvious from this investigation, however, that more information is sorely needed to strengthen the foundation on which health policy on screening in the elderly can be established. ☐

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APPENDIX A—References Reviewed but Not Abstracted

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The Increasing Supply of Physicians in US Urban and Rural Areas, 1975 to 1988

Paul D. Frenzen, PhD

ABSTRACT

Background. Despite the rapid growth of the US physician supply since the mid-1970s, it remains unclear whether physicians have spread into the most rural areas of the country. This report examines the urban–rural distribution of physicians between 1975 and 1988.

Methods. A county-based typology of the urban–rural continuum was employed to examine trends in the supply of nonfederal primary care physicians, specialist physicians, and osteopaths.

Results. All urban and rural areas gained physicians during the late 1970s and 1980s. The supply of physicians increased most rapidly in metropolitan counties. Within nonmetropolitan areas, urbanized remote counties became more prominent centers of the physician supply. Osteopaths were more likely to locate in the most rural areas than allopaths. The physician supply in all areas also became more specialized over time.

Conclusions. The rapid growth of the US physician supply was associated with the spread of more practitioners into all parts of the country. However, the supply of physicians increased most rapidly in urban areas, widening urban–rural differences in the availability of physicians. (*Am J Public Health.* 1991; 81:1141–1147)

Introduction

The number of nonfederal allopathic and osteopathic physicians in the United States increased nearly four times more rapidly than the population between 1975 and 1988. Despite the rapid increase, physicians remained concentrated in urban centers. By 1988 the ratio of physicians per 100 000 persons was 226 in metropolitan areas and only 98 in nonmetropolitan areas. This gap has persisted despite government efforts to place more physicians in rural communities and concerns about an imminent national surplus of physicians.

The uneven distribution of physicians is partly due to the concentration of medical education, research, and specialized care in urban central places. However, there is much evidence that physicians prefer urban locations for other reasons, including the personal ties developed during medical training, the greater opportunities to interact with professional colleagues, and a taste for urban amenities.[1] The disproportionate tendency of physicians to locate in urban areas narrows the choice of practitioners for rural populations and suggests that physician services are relatively less accessible in rural areas.

The federal government first began to address the geographic maldistribution of physicians during the early 1970s, when a variety of measures were enacted to increase the supply of practitioners in inner cities and rural areas.[2,3] One of the most important measures was the National Health Service Corps (NHSC) scholarship program, which provided financial aid to medical students and other health trainees in return for service after graduation in designated Health Personnel Shortage Areas (HPSAs, formerly termed Health Manpower Shortage Areas).

Federal efforts to increase the number of physicians in underserved areas were curtailed during the early 1980s, and the NHSC scholarship program was gradually phased out.[2,4] The shift in policy was influenced by the 1980 report by the Graduate Medical Education National Advisory Committee (GMENAC), which predicted that a substantial surplus of physicians would emerge by 1990.[5] Federal officials also contended that market forces alone would be sufficient to attract more physicians to rural areas as competition intensified among physicians in urban centers.[2]

A series of studies by the RAND Corporation during the early 1980s provided evidence that physicians' location decisions had been affected by market forces.[6–10] The studies showed that the proportion of small towns served by various kinds of specialists had increased during the 1970s until nearly every urban place with 2500 or more inhabitants had at least one physician. The RAND researchers argued that physicians had responded to the increasing physician supply in accordance with economic location theory by diffusing to smaller places until all towns above a critical size had at least one physician. They also predicted that physicians would spread more widely in rural

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Paul D. Frenzen is with the US Department of Agriculture in Washington, DC. Requests for reprints should be sent to Paul D. Frenzen, PhD, Economic Research Service, US Department of Agriculture, 1301 New York Avenue, NW, Room 340, Washington, DC 20005.

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areas during the 1980s as the growth of the physician population continued.[7]

In contrast, trends in physician/population ratios during the 1970s and early 1980s led other observers to conclude that physicians had not spread into rural areas in substantial numbers.[11–14] Metropolitan (metro) counties continued to have much higher physician/population ratios than nonmetropolitan (nonmetro) counties. Thus, physicians remained concentrated in urban centers. Further classification of counties by population size also showed that physican/population ratios had increased most slowly in small nonmetro counties. This trend was interpreted as evidence that few physicians had moved into the most rural areas of the country.[11,13]

Continued public concern about physician shortages led to new federal initiatives to place physicians in underserved communities during the late 1980s, including the NHSC loan repayment program and a Medicare reimbursement bonus for services performed by physicians in designated HPSAs.[4] Funding for the NHSC scholarship program was also significantly increased starting in 1991.[15] The new initiatives reveal that some areas remained unattractive to physicians despite the rapid growth of the physician supply.

The purpose of this report is to examine the growth of the nonfederal physician supply since the mid-1970s during a period of major changes in federal health personnel policy. A county-based typology of the urban–rural continuum will be employed to determine whether the supply of physicians increased in predominantly rural areas without sizeable towns, as well as in urbanized rural areas and smaller and larger metro centers. These distinctions are important because the RAND studies demonstrated that town size plays a critical role in physicians' location decisions. Other studies of the physician supply based on the classification of counties by population size have failed to identify predominantly rural areas because nonmetro counties with small populations may lie near metro centers or contain relatively large urban places.

Methods

Information about the annual distribution of patient care allopaths (MD physicians) and active osteopaths (DO physicians) by county was obtained from the 1989 Area Resource File and the 1975 to 1988 MD/DO Enhanced County File prepared by the Office of Data Analysis and Management, US Department of Health and Human Services. The data for allopaths and osteopaths were not always available for the same year. Federal physicians were excluded because most served special populations. Alaska was also excluded because the physician data for the state were not broken down by county. Additional information about the distribution of the resident population by county was obtained from the US Bureau of the Census intercensal population estimates. Counties affected by border changes between 1975 and 1988 were combined into larger units with constant boundaries.

Nonfederal allopaths were divided into two groups according to self-designated practice specialty: (1) primary care physicians providing patient care in general/family practice, internal medicine, obstetrics/gynecology, or pediatrics; and (2) specialist physicians providing patient care in any other specialty. Nonfederal active osteopaths were combined in a third category. Resident physicians were included in the appropriate category because they provided a substantial amount of patient care. Internal medicine subspecialist physicians were placed in the primary care category because they could not be distinguished from general internists prior to 1986. The small proportion of physicians who did not report a specialty were excluded from the analysis. Active osteopaths included both patient care and non–patient care physicians because the two groups could not be distinguished after 1981, when nearly 95% were engaged in patient care.

All counties were classified into 10 urban–rural continuum categories by using a typology developed by the Economic Research Service, US Department of Agriculture.[16] Counties were first divided into metro and nonmetro areas by using the 1983 US Office of Management and Budget definition of Metropolitan Statistical Areas (MSAs). Metro counties were then subdivided by 1980 MSA population and structure, and nonmetro counties were subdivided by 1980 county urban population and proximity to MSAs (Table 1). Variations in the level of urban influence are thus measured by the population and structure of MSAs in metro areas and by county urban populations and proximity to MSAs in nonmetro areas. Counties toward the urban end of the continuum are described as "more urban" and those toward the rural end are described as "more rural" for convenience.

The supply of physicians in each category was measured by the ratio of physicians per 100 000 residents. Separate ratios were calculated for primary care and specialist physicians for 1975, 1981, and 1988, and for osteopaths for 1976, 1981, and 1987. Since the reporting date for physicians shifted from December 31 to January 1 in 1988, the 1988 ratios were based on the 1987 resident population to increase comparability over time. The 1981 to 1988 period was only 1 day longer than the 1975 to 1981 period because of the shift in dates.

Changes in the physician supply over time were measured by calculating the mean annual change in the physician/population ratios for each urban–rural continuum category. This measure of absolute change was chosen to compare areas with different initial ratios.

Results

Primary Care Physicians

The total number of primary care physicians grew 3.8% per year between 1975 and 1988, rising from 132 000 to 207 700. During the same period the overall ratio of primary care physicians per 100 000 persons increased from 61.4 to 85.5 (Table 2). However, the ratio was substantially higher at the more urban end of the urban–rural continuum than at the more rural end (Figure 1). The central areas of large MSAs (large metro core counties) had the highest ratio throughout the period. The suburbs of large MSAs (large metro fringe counties) and medium metro counties had the next highest ratios. Small metro counties had lower ratios than other metro counties, but higher ratios than nonmetro counties.

Within nonmetro areas, urbanized counties had higher primary care ratios than less urbanized or rural counties. However, within each urban population category, remote counties some distance from MSAs had more primary care physicians than adjacent counties near MSAs. This pattern indicates that nonmetro primary care physicians preferred to locate in areas where distance made metro physician services less accessible to local residents.

The mean annual change in the overall primary care physician ratio was nearly constant, varying from 1.98 in 1975 to 1981 to 2.04 in 1981 to 1988 (Table 2). On average, communities thus gained about two primary care physicians per 100 000 residents per year during the late 1970s and 1980s. Although all urban and rural areas gained primary care physicians, the annual gains were greatest in metro counties and nonmetro urbanized remote counties. All other nonmetro counties added primary care physicians much more slowly.

The faster growth of the primary care physician supply in more urban areas caused urban–rural differences in the availability of practitioners to widen over time. The gap between metro and nonmetro areas expanded from 30 to 45 primary care physicians per 100 000 persons between 1975 and 1988 (Table 2). Differences within nonmetro areas also increased as nonmetro urbanized remote counties (and to a lesser extent, less urbanized remote counties) became more prominent centers of the primary care physician supply (Figure 1). The emergence of these centers indicates that nonmetro primary care physicians have become increasingly concentrated in counties with larger urban places, even though all areas gained practitioners.

Nonmetro counties gained primary care physicians more rapidly during the 1980s than the late 1970s (Table 2). In contrast, most metro counties (except large metro fringe counties) gained primary care physicians more slowly. The shift in growth rates indicates that nonmetro areas and the suburbs of large MSAs became more attractive to primary care physicians after 1981.

The urban–rural differences in the supply of primary care physicians were affected by the classification of internal medicine subspecialists as primary care physicians. However, subspecialists accounted for only a small part of the differences in 1988 when subspecialists could be distinguished from general internists. In that year 8% of the primary care physicians in metro areas and 3% of those in nonmetro areas were actually subspecialists.

In contrast, resident physicians accounted for a larger share of the urban–rural differences in the primary care physician supply. More than 20% of the primary care physicians in metro areas were in residency training in 1988, in comparison with only 4% of those in nonmetro areas. There was little change in these proportions over time. The differences in the availability of primary care physicians would be smaller if residents were excluded, but the important role of residents in staffing urban teaching hospitals would be ignored.

TABLE 1—County Urban–Rural Continuum Categories

Urban–Rural Continuum Category	Number of Counties	Percent of 1987 Resident Population
Total counties[a]	3070	100.0
Metro counties	710	76.8
1. Large metro counties		
a. Core counties containing primary central city of MSAs with 1980 populations of 1 000 000 or more	55	29.0
b. Fringe (suburban) counties of MSAs with 1980 populations of 1 000 000 or more	169	16.3
2. Medium metro counties in MSAs with 1980 populations of 250 000 to 999 999	289	22.4
3. Small metro counties in MSAs with 1980 populations under 250 000	197	9.1
Nonmetro Counties	2360	23.2
1. Urbanized counties with 20 000 or more urban residents in 1980		
a. Adjacent to an MSA[b]	137	4.4
b. Remote (not adjacent to an MSA)	150	3.7
2. Less urbanized counties with 2500 to 19 999 urban residents in 1980		
a. Adjacent to an MSA[b]	551	6.2
b. Remote (not adjacent to an MSA)	748	6.3
3. Rural counties with fewer than 2500 urban residents in 1980		
a. Adjacent to an MSA[b]	228	1.1
b. Remote (not adjacent to an MSA)	546	1.6

[a] Excluding Alaska.

[b] Counties were considered adjacent to an MSA when the county and MSA borders touched at more than one point and 2% or more of the county labor force commuted to the central county of the MSA.

Note: Metro = metropolitan; MSA = Metropolitan Statistical Area.

TABLE 2—Nonfederal Primary Care Physicians per 100 000 Resident Population by Urban–Rural Continuum Category, 1975–1988[a]

Urban–Rural Continuum Category	Primary Care Physician/ Population Ratio			Mean Annual Change in Primary Care Physician/ Population Ratio	
	1975	1981	1988	1975–1981	1981–1988
US Total	61.4	73.3	85.5	1.98	2.04
Metro	68.5	82.3	95.9	2.31	2.26
Large metro					
Core	86.6	101.9	116.0	2.56	2.34
Fringe	55.9	69.6	85.7	2.28	2.68
Medium metro	59.3	72.9	85.1	2.28	2.02
Small metro	52.0	65.1	76.2	2.18	1.86
Nonmetro	38.6	44.4	51.3	0.97	1.14
Urbanized					
Adjacent to MSA	42.1	47.9	54.8	0.98	1.14
Remote	46.3	57.8	69.3	1.91	1.92
Less urbanized					
Adjacent to MSA	35.4	39.1	44.2	0.61	0.86
Remote	39.1	45.0	51.8	0.98	1.13
Rural					
Adjacent to MSA	28.1	29.0	31.9	0.16	0.48
Remote	29.0	32.8	38.2	0.63	0.91

[a] Excluding Alaska.

Note: Metro = metropolitan; MSA = Metropolitan Statistical Area.

Specialist Physicians

During 1975 to 1988 the number of specialist physicians grew more rapidly than the number of primary care physi-

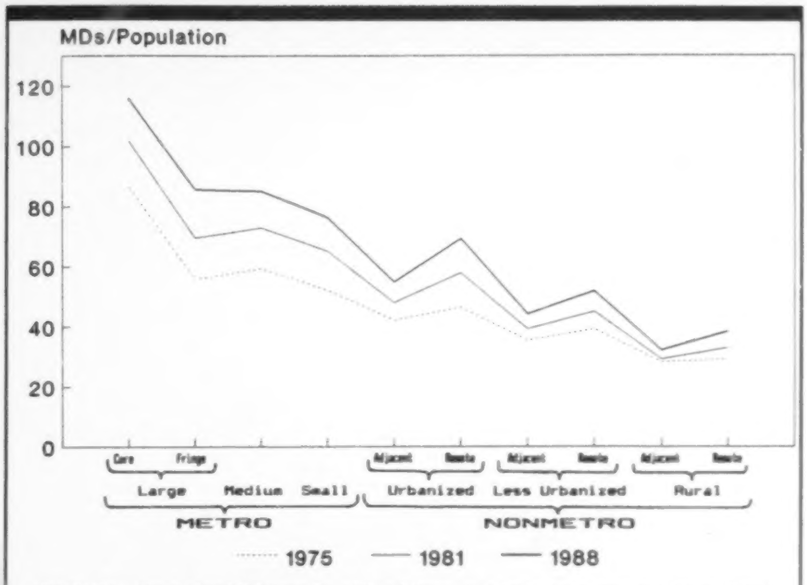

FIGURE 1—Nonfederal Primary Care Physicians per 100 000 persons by Urban–Rural Continuum Category, 1975–1988

cians, increasing by 4.2% per year from 147 200 to 242 900. As a result, the overall specialist physician population ratio rose from 68.4 to 100.0 (Table 3). The urban–rural distributions of specialists and primary care physicians were very similar (Figure 2). However, there were even greater differences between more urban and more rural areas in the supply of specialists. The specialist ratio was highest in the central areas of large MSAs and lowest in nonmetro rural counties. Within nonmetro areas, remote counties had more specialists than adjacent counties in the same urban population category.

The total supply of specialists grew more rapidly during the 1980s than the late 1970s (Table 3). All urban and rural areas gained specialists, but the gains were greater in more urban areas, repeating the pattern among primary care physicians. The increase in the specialist supply was particularly rapid in large MSAs, which gained more than three specialists per 100 000 persons per year. In contrast, the specialist supply grew only one tenth as rapidly in nonmetro rural counties.

TABLE 3—Nonfederal Specialist Physicians per 100 000 Resident Population by Urban–Rural Continuum Category, 1975–1988*

Urban–Rural Continuum Category	Specialist Physician/ Population Ratio			Mean Annual Change in Specialist Physician/ Population Ratio	
	1975	1981	1988	1975–1981	1981–1988
US Total	68.4	83.6	100.0	2.53	2.74
Metro	81.7	99.3	118.2	2.94	3.14
Large metro					
Core	105.1	124.3	145.7	3.21	3.56
Fringe	63.8	83.4	103.0	3.26	3.27
Medium metro	69.8	86.4	102.9	2.77	2.75
Small metro	63.5	79.4	95.4	2.65	2.66
Nonmetro	25.8	33.4	39.9	1.26	1.07
Urbanized					
Adjacent to MSA	36.2	46.6	54.7	1.73	1.36
Remote	50.7	64.9	77.5	2.37	2.10
Less urbanized					
Adjacent to MSA	16.9	22.0	26.5	0.84	0.76
Remote	21.2	27.7	32.9	1.09	0.87
Rural					
Adjacent to MSA	5.8	7.6	10.1	0.30	0.41
Remote	6.3	9.3	11.0	0.49	0.29

*Excluding Alaska.
Note: Metro = metropolitan; MSA = Metropolitan Statistical Area.

Urban–rural differences in the availability of specialists widened over time due to the faster growth of the specialist ratio in more urban areas (Table 3). Between 1975 and 1988 the gap between metro and nonmetro areas expanded from 56 to 78 specialists per 100 000 persons. In nonmetro areas, urbanized remote counties also became more prominent centers of the specialist supply (Figure 2). Most nonmetro areas (except rural adjacent counties) gained specialists more slowly after 1981, but the central areas of large MSAs began gaining specialists even more rapidly.

The classification of internal medicine subspecialists as primary care physicians reduced urban–rural differences in the specialist ratio (see above). However, the inclusion of specialist residents increased the differences. During the late 1970s and 1980s about 17% of the specialists in metro areas were in residency training, in contrast to only 3% of those in nonmetro areas. The inclusion of residents is appropriate in view of their role in urban teaching hospitals.

Osteopaths

The number of osteopaths also grew rapidly between 1976 and 1987, increasing by 4.2% per year from 13 700 to 21 800. The osteopath/population ratio rose from 6.3 to 9.0 during the same period (Table 4). Urban–rural differences in the supply of osteopaths were small in comparison with the differences in the supply of allopaths. However, osteopaths were distributed across the urban–rural continuum in a distinct pattern, with the greatest supply concentrated in the most urban and most rural areas of the country (Figure 3).

The supply of osteopaths increased more rapidly after 1981 (Table 4). Some areas (nonmetro urbanized remote and less urbanized adjacent counties) actually lost osteopaths during the late 1970s. In contrast, all areas gained osteopaths after 1981, with the fastest growth occurring in the most urban and most rural areas.

Composition of the Physician Supply

The uneven distribution of primary care physicians, specialist physicians, and osteopaths across the urban–rural contin-

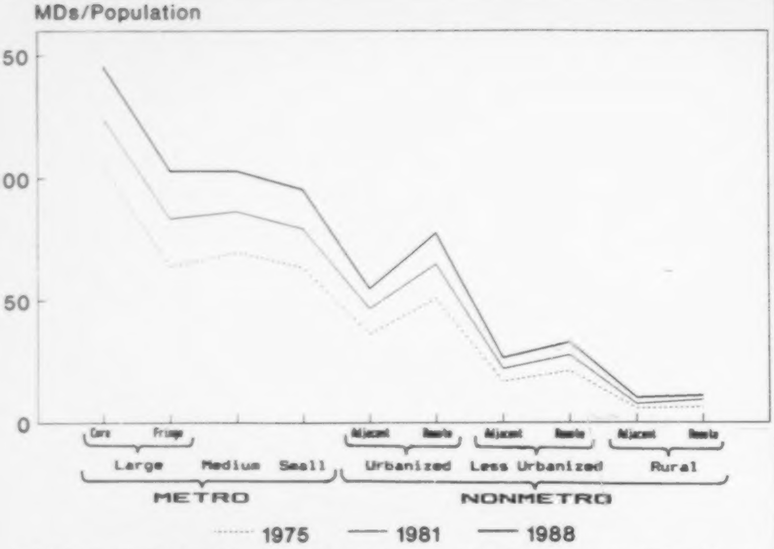

FIGURE 2—Nonfederal Specialist Physicians per 100 000 Persons by Urban–Rural Continuum Category, 1975–1988

uum produced major differences in the composition of the physician supply (Table 5). Primary care physicians represented a larger share of the physician supply in more rural areas, accounting for more than two thirds of the practitioners in nonmetro rural counties. In contrast, the proportion of specialist physicians was greater in more urban areas. The higher degree of physician specialization in more urban areas reflected the concentration of medical referral centers in cities and larger towns.

Osteopaths represented less than 5% of the physician supply during the late 1970s and 1980s (Table 5). However, they provided a disproportionate share of the physicians in more rural areas because they were relatively more willing to locate in nonmetro, less urbanized and rural counties than allopaths (Figures 1–3). By 1987 to 1988 more than 10% of the physicians in nonmetro rural counties were osteopaths. The role of osteopaths was even greater in eight states (Arizona, Florida, Michigan, Missouri, New Jersey, Ohio, Pennsylvania, and Texas) where licensing legislation was historically most favorable to the osteopathic profession.[17] In these states osteopaths accounted for 36% of the physicians in nonmetro rural counties.

The increasing specialization of the physician population had a direct impact on the composition of the physician supply. From 1975 to 1976 and 1987 to 1988 the proportion of primary care physicians declined, and the proportion of specialist physicians increased in every category of the urban–rural continuum (Table 5). The increase in the proportion of specialists was greatest in nonmetro counties.

Discussion

The evidence examined here indicates that the supply of nonfederal physicians increased in every part of the country. By the late 1980s all areas had more primary care and specialist physicians and osteopaths than in the mid-1970s. The rapid growth of the US physician supply was thus associated with the movement of more practitioners into predominantly rural areas, as well as urbanized rural and urban areas. However, the physician supply increased more rapidly in metro counties and nonmetro urbanized remote counties than in other nonmetro areas. As a result, urban–rural differences in the availability of physicians widened over time.

A major reason for the slow growth of the physician supply in predominantly rural areas may have been the limited number of places large enough to support physicians. Only 17% of the population of nonmetro rural counties lived in places with 1000 or more inhabitants in 1980.[18] The remainder of the population was dispersed in communities seemingly too small to support physicians, despite the spread of practitioners into more rural areas. However, this analysis was not designed to reveal the size of places within counties that actually gained physicians.

Osteopaths played a surprisingly important role in predominantly rural areas because they were relatively more likely to locate there than allopaths. As a result, osteopaths accounted for a disproportionate share of the physician supply in nonmetro rural counties, particularly in the eight states historically most hospitable to the profession. The factors that account for the greater willingness of osteopaths to locate in predominantly rural areas de-

TABLE 4—Nonfederal Active Osteopaths per 100 000 Resident Population by Urban–Rural Continuum Category, 1976–1987[a]

Urban–Rural Continuum Category	Active Osteopaths/Population Ratio			Mean Annual Change in Active Osteopath/Population Ratio	
	1976	1981	1987	1976–1981	1981–1987
US Total	6.3	7.0	9.0	0.13	0.33
Metro	6.9	7.7	9.8	0.17	0.35
Large metro					
Core	7.4	8.3	10.7	0.19	0.40
Fringe	7.7	8.6	10.5	0.17	0.32
Medium metro	6.5	7.5	9.6	0.19	0.35
Small metro	4.5	4.6	6.1	0.02	0.25
Nonmetro	4.6	4.7	6.2	0.03	0.24
Urbanized					
Adjacent to MSA	4.3	4.7	6.2	0.09	0.25
Remote	3.5	3.4	4.9	−0.01	0.25
Less urbanized					
Adjacent to MSA	4.7	4.7	5.7	0.00	0.16
Remote	5.2	5.4	7.0	0.04	0.27
Rural					
Adjacent to MSA	3.8	4.0	4.8	0.05	0.13
Remote	5.6	5.6	8.3	0.00	0.44

[a] Excluding Alaska.

Note: Metro = metropolitan; MSA = Metropolitan Statistical Area.

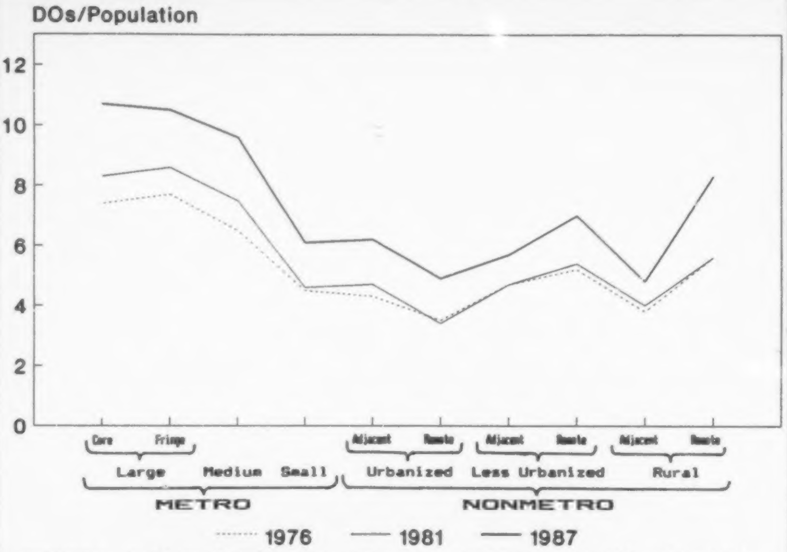

FIGURE 3—Nonfederal Active Osteopaths per 100 000 Persons by Urban–Rural Continuum Category, 1975–1988

serve further investigation in view of the reluctance of allopaths to practice in the same locations.

Despite the slow growth of the physician supply in predominantly rural areas, all nonmetro areas gained primary care physicians more rapidly during the 1980s than before. The faster growth may have been due to two different factors. First, increasing competition among physicians in metro areas may have encouraged more practitioners to select nonmetro locations. The slower growth of the primary care physician supply in most metro areas after 1981 supports this interpretation. Second, the number of general practice physicians decreased more slowly during the 1980s than the 1970s, elevating the net growth of the primary care physician supply.[19] The impact was greatest in nonmetro areas, where general practitioners represented a larger share of primary care physicians.

TABLE 5—Percent Composition of the Nonfederal Physician Population by Urban–Rural Continuum Category, 1975–1976 and 1987–1988[a]

	Percent of Nonfederal Physician Population[b]					
	1975–1976			1987–1988		
Urban–Rural Continuum Category	Primary Care Physicians	Specialist Physicians	Active Osteopaths	Primary Care Physicians	Specialist Physicians	Active Osteopaths
	(%)	(%)	(%)	(%)	(%)	(%)
US Total	45.1	50.2	4.7	44.0	51.4	4.6
Metro	43.6	52.0	4.4	42.8	52.8	4.4
Large metro						
Core	43.5	52.8	3.7	42.6	53.5	3.9
Fringe	43.8	50.0	6.1	43.0	51.7	5.3
Medium metro	43.7	51.4	4.9	43.1	52.1	4.9
Small metro	43.3	52.9	3.8	42.9	53.7	3.4
Nonmetro	55.9	37.3	6.8	52.7	41.0	6.3
Urbanized						
Adjacent to MSA	50.9	43.8	5.3	47.3	47.3	5.4
Remote	46.0	50.4	3.5	45.7	51.1	3.2
Less urbanized						
Adjacent to MSA	62.0	29.6	8.4	57.9	34.7	7.4
Remote	59.6	32.3	8.1	56.4	35.9	7.7
Rural						
Adjacent to MSA	74.5	15.4	10.2	68.2	21.5	10.3
Remote	70.7	15.4	13.9	66.4	19.2	14.4

[a] Excluding Alaska.
[b] Percents may not sum to 100.0% because of rounding.
Note: Metro = metropolitan; MSA = Metropolitan Statistical Area.

Specialist physicians in nonmetro areas were affected by different factors than primary care physicians because most nonmetro areas gained specialists more slowly after 1981. One possible factor may have been the limited number of places large enough to support specialists. By 1980 more than 36 million nonmetro residents lived outside places with 2500 or more persons in communities that the RAND studies showed rarely had specialists. However, more than 86% of this population lived in open country or places with fewer than 1000 persons that may never be able to support specialists.[18]

During the late 1970s and 1980s the physician supply in all areas also became more specialized. This trend was particularly dramatic in nonmetro rural counties, where the number of primary care physicians per specialist dropped from nearly five in 1975 to slightly more than three in 1988. The impact on the supply of medical services was unclear because specialists in rural areas may have spent much of their time providing primary care. The growth in the number of rural specialists also raised questions about the size of the case load needed to maintain physician expertise in specialized diagnostic and therapeutic procedures.[10]

Another important trend in nonmetro areas was the growing prominence of nonmetro urbanized remote counties as local centers of the physician supply. The rapid growth of the physician/population ratio and the higher proportion of specialists in these counties reflected the continuing development of secondary medical centers providing a variety of health services for rural areas some distance from major cities.[20] These centers were distinct from the tertiary medical centers in metro areas providing more specialized services for larger regions. The closing of hospitals in smaller rural communities and the greater competition from metro physicians in areas closer to cities may have contributed to the increasing concentration of physicians in rural secondary medical centers.

Nonmetro, less urbanized remote counties also became local centers of the primary care physician supply. These centers had fewer primary care physicians than nonmetro urbanized remote counties and did not have many specialist physicians. As a result, they appeared to be rural primary medical centers serving areas some distance from larger towns.

The most striking trend in metro areas was the continued rapid growth in the physician supply. It remains unclear whether this trend has produced a physician surplus because the demand for physician services depends on a complex variety of factors, including the costs of medical care, changes in health insurance coverage, advances in medical technology, the supply of midlevel health practitioners, and shifts in health care needs.

The even more rapid growth of the physician supply in large MSAs also deserved attention. By 1988 the suburbs of large MSAs had more primary care and specialist MDs than medium or small MSAs, reversing earlier patterns. The continued rapid growth was possible in part because physicians in large MSAs spent less time in patient care activities than those in other locations, restraining the actual growth in the supply of physician services.[22]

The variations in physician–population ratios described in this report do not necessarily reflect unequal access to physician services. Because many rural residents travel to urban areas to visit physicians, the average patient load per physician in rural areas is smaller than physician–population ratios suggest.[21] Rural physicians also work longer hours and see more patients than urban physicians, and this pattern reduces urban–rural differences in the supply of physician services.[22] Nevertheless, access to physician services outside densely populated urban areas still depends to a considerable extent on the location of physicians. In addition, rural communities have higher proportions of the poor and elderly than urban areas and may therefore require more physicians to meet local medical needs.[23]

Programs to provide more physician services in underserved areas need to take account of the fact that the most rural areas of the country remain unattractive to physicians. The present distribution of physicians in nonmetro areas suggests that it may be worth considering measures to encourage practitioners already located in rural primary and secondary medical centers to provide more care in surrounding areas. Some of the mechanisms that might be used to expand medical services into surrounding areas from these centers include mobile and satellite clinics, an increased role for midlevel health practitioners, and special transportation arrangements for nonemergency patients. ☐

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Cognitive Impairment and Hospital Use

ABSTRACT

Background. An increasing number of older people are at higher risk of developing Alzheimer's Disease or another dementias. The resultant cognitive impairment has been well identified as one of the risk factors for nursing home placement but it has not been well studied as a risk for hospitalization.

Methods. To study the association between cognitive impairment and hospital use, this study examined data from a randomly selected cohort of community dwelling California elderly (California Senior Survey). The cohort was followed for one year (N = 940).

Results. Only about 17 percent of the study population was in the mildly or moderately/severe impaired category. Those participants with cognitive impairment were almost twice as likely as those without any impairment to be hospitalized and to be hospitalized for six days or more.

Conclusions. Although the cause of the cognitive impairment and hospitalization was not determined in this study, the data support the importance of cognitive impairment as a risk factor for hospital use even after controlling for age, ADL, IADL, and prior hospitalizations. The causes for this need to be further examined so the burden of such care can be reduced. (*Am J Public Health* 1991;81:1153–1157.)

Philip G. Weiler, MD, MPH, James E. Lubben, MPH, DSW, and Iris Chi, DSW*

Introduction

Cognitive status has been identified as one of the risk factors for nursing home placement[1] but it has not been well studied as a cause of hospitalization. Although several studies have looked at health services utilization due to Alzheimer's Disease or other dementias,[2–4] these studies have been mainly interested in estimating costs of care, particularly cost of nursing home care.[4] Further, they have relied largely on cross-sectional data sources and the number of individuals with cognitive impairment in some of these data sets has been small.

Few studies have looked at the hospitalization admission rate and length of stay in a random sample of community dwelling elderly based on their cognitive status. Huang and Hu[2] found that the elderly demented utilized an average of 4.81 days of short-term acute hospital care per person. This represented an additional utilization of short stay hospitals due to Alzheimer's Disease of 0.81–1.02 days per year based on the national hospital discharge survey data. Coughlin and Liu,[5] using data from the 1981–82 National Long-Term Care Channeling Demonstration Project and the Short Portable Mental Status Questionnaire (SPMSQ) to rate a person's cognitive status, found that the severely cognitively impaired spent a higher proportion of survival days in hospitals (.10) then mild/no cognitive impaired (.08) after six months of follow-up. More recently, Binder and Robins[6] found that community dwelling elderly were more likely to be hospitalized if they scored poorly on the Folstein Mini-Mental State Exam.

In response to the limited research on this topic, the purpose of the present study was to examine the relation between cognitive status and hospitalization in a randomly selected cohort of community dwelling Medicaid recipients. Additionally, the study examined the relation of cognitive status and length of hospital stays. The cohort was followed for one year.

Methods

Design

The data were drawn from the California Senior Survey (CSS), a random sample of community dwelling elderly MediCal (Medicaid) recipients from eight representative communities within California. The CSS respondents were randomly selected from files which contained the names, addresses, and other basic demographic data with all MediCal recipients. Only MediCal recipients 65 years of age and older were sampled. CSS respondents were interviewed face to face in their homes in 1982. Comparisons between the survey data and statewide data indicated that respondents were typical of California's MediCal elderly population.[7] More complete descriptions of the sampling methodology and sample characteristics have already been published.[7,8]

Hospital utilization data were obtained for one year following the inter-

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*Dr. Philip G. Weiler has died since this study was completed. Dr. Lubben is Associate Professor, School of Social Welfare, UCLA; Dr. Chi is Lecturer, Department of Social Work, University of Hong Kong. Address reprint requests to Dr. James Lubben, UCLA, School of Social Welfare, 247 Dodd Hall, 405 Hilgard Ave, Los Angeles, CA 90024. This paper, submitted to the Journal December 19, 1989, was revised and accepted for publication January 7, 1991.

TABLE 1—Selected Characteristics of Study Cohort by Level of Cognitive Impairment (percentage distribution) N = 940

	Not Impaired	Mildly Impaired	Mod/Sev Impaired
(Total %)	(8[illegible]%)	(9.3%)	(7.4%)
Sex			
Males	89.3	5.4	5.3
Females	81.0	10.8	8.3
Age (years)			
65–74	91.1	6.1	2.8
75–84	81.2	10.1	8.7
85+	67.2	15.7	17.2
Marital Status			
Married	82.9	8.3	8.8
Widowed	81.4	11.0	7.6
Separated/Divorced	88.7	6.4	5.0
Single	89.1	3.6	7.3
Living Arrangement			
Alone	87.6	8.5	4.0
With Others	78.0	10.4	11.6
Race/Ethnicity			
White	86.3	7.4	6.3
Hispanic	78.7	13.1	8.2
Black	79.1	10.8	10.1
Asian	79.4	11.2	9.4
Education			
None	51.7	20.0	28.3
Elementary	83.3	9.3	7.3
High School	88.6	5.9	5.6
Beyond High School	84.8	11.2	3.9

view. The hospital utilization data came from MediCal paid claims and computer tapes which were merged with data obtained from Medicare intermediaries.[9] Combining these two sources provided data on number of days hospitalized for 92 percent of the original CSS sample (N = 1,037). To assure accuracy, the computer tapes were examined for any bills submitted for hospitalization during the follow-up year to either MediCal or Medicare intermediaries on behalf of CSS respondents. To overcome any problems that may have been caused by late billing, bills submitted up to two years following the interview were reviewed.

Besides those without hospital data, 15 other respondents were also dropped from the present study. Six lacked cognitive status data and nine had died within a few months of the survey and so their hospital data were necessarily censored. Analyses were run with and without these six respondents and they did not significantly affect the outcome. However, dropping these six respondents from the analyses was considered the more appropriate method. Although 25 other respondents also had died within a year of the baseline survey, they all had extensive hospital stays before their death and so it was possible to properly classify their hospital use for the present analyses.

No significant differences in terms of major sociodemographic and health status characteristics were noted between survey respondents with and without complete data. All respondents with complete hospital and cognitive status data were used for the study analyses (N = 940).

A structured measure of mental status was administered, the Short Portable Mental Status Questionnaire (SPMSQ). This instrument has high validity and test-retest reliability.[10] The widely used test is a 10-item questionnaire that is designed to measure several areas of intellectual domain including orientation, short- and long-term memory, general information, and problem solving. Pfeiffer[10] originally recommended adjustments for race and education. However, more recent versions of the SPMSQ only adjust for education.[11,12] Thus the present analysis chose to only adjust SPMSQ scores for education, although both race and education were entered as covariates in the multiple regression analysis.

Using Pfeiffer's recommendations, the SPMSQ scores used to categorize participants were: not impaired (0–2 errors), mildly impaired (3–4 errors), and moderately to severely impaired (over 4 errors).[10] The SPMSQ was administered in the elderly person's home by trained interviewers who also administered a comprehensive questionnaire composed of several previously described instruments.[13] Included in this battery were two commonly used functional status scales: the Activity of Daily Living (ADL) and Instrumental Activity of Daily Living (IADL) scale.[14,15]

Two dichotomous variables measuring hospital utilization were used: whether a respondent had been hospitalized within the year following the interview, and whether a respondent experienced a hospital stay of six or more days in the year following the interview. The figure of six days was selected because it was slightly higher than the average length of stay for all respondents (mean length of stay = 5.2 days).

Analysis

The present analyses used bivariate analyses (Chi-square) and odds ratios. A series of logistic regression models were used to examine the relation of the levels of cognitive status with the two indices of subsequent hospital utilization. Age, education, race, and prior hospitalizations during the year prior to the survey were controlled. The statistical program used was SAS and analyses were conducted on an IBM 3090 mainframe computer.

Study Population

Reflecting the diversity of California's ethnic population, the sample included a sizable portion of Blacks (17 percent), Hispanic (13 percent), and Asians (12 percent), although a majority (58 percent) were White. Seventy-two percent of the sample were females and over 60 percent of the sample lived alone. The age distribution of the sample was 42 percent (65–74 years of age), 44 percent (75–84 years old), and 14 percent (85 years old and over). The majority of participants were widowed (56 percent) with 23 percent being married, 6 percent single, and 15 percent separated or divorced. The educational level obtained by the participants was 33 percent finished high school and 19 percent had an education beyond high school, 6 percent had no formal education, and about 42 percent finished elementary school.

Results

Table 1 gives selected characteristics of the participants by level of cognitive impairment. Overall about 17 percent of the study population was either mildly or moderately to severely impaired. There was a higher prevalence of cognitive impairment in females than males. The prevalence of cognitive impairment increased significantly with age. Marital status was not related to cognitive impairment but cognitive impairment was more prevalent in those living with others than in those living alone. There were no race or ethnicity differences in the prevalence of cognitive impairment, but lack of education was associated with cognitive impairment.

TABLE 2—Self-care Capacity of Study Cohort by Level of Cognitive Impairment (percentage distribution) N = 940

	Not Impaired	Mildly Impaired	Mod/Sev Impaired
(Total %)	(83.3%)	(9.3%)	(7.4%)
ADL Dependencies			
0	88.2	7.4	4.4
1	85.1	8.9	6.1
2+	61.1	16.8	22.1
IADL Dependencies			
0	92.5	6.5	0.9
1	95.2	3.7	1.1
2	88.3	7.8	3.9
3	80.3	12.2	7.5
4	82.4	9.8	7.8
5+	55.4	18.9	25.8

TABLE 3—Co-morbidity of Study Cohort by Level of Cognitive Impairment (percentage distribution) N = 940

	Not Impaired	Mildly Impaired	Mod/Sev Impaired
(Total %)	(83.3%)	(9.3%)	(7.4%)
Chronic Conditions			
Arthritis	86.3	7.4	6.3
Stroke	69.4	10.2	20.4
High Blood Pressure	84.9	8.4	6.7
Heart Condition	81.6	9.9	8.5
Respiratory	87.3	6.4	6.4
Gastrointestinal	85.6	7.1	7.4
Diabetes	82.9	8.6	8.6
Number of Conditions			
0	76.3	12.5	11.3
1	81.3	13.7	5.1
2	87.0	6.8	6.3
3	85.8	7.8	6.5
4+	81.8	8.5	9.7

There was a significantly higher prevalence of cognitive impairment in those participants with two or more ADL dependencies. Similarly, an increased number of instrumental activities of daily living (IADL) dependencies was also significantly associated with cognitive impairment with money management having the strongest association (Table 2). The number of total chronic conditions was not significantly related to cognitive impairment (Table 3). Arthritis was negatively associated with cognitive impairment whereas stroke was positively related.

Odds ratios adjusted for age, education, race, and any hospitalizations during year prior to survey are shown in Table 4. Those participants with cognitive impairment were more likely to be hospitalized and to have increasing length of hospitalizations. Similarly, greater impairments on the ADL and IADL scales were associated with increased risk of hospitalization.

The number of dependencies on the ADL and IADL was collapsed into categories to facilitate analyses. Very few respondents had more than two ADL dependencies and so a single dependency on the ADL was considered a mild ADL impairment whereas being unable to perform two or more ADL functions was considered to be a moderate to severe ADL impairment. Cut points for the IADL were selected to maximize the relationship between IADL and hospital use. One to two IADL dependencies were judged to be a mild IADL impairment, three to four a moderate impairment, and five or more IADL dependencies were classified as a severe dependency. Any level of impairment on the SPMSQ was predictive of hospital use. Increased impairment on any of the scales was associated with increased risk of extended hospital stays. Further, Table 4 illustrates that those elderly with severe IADL impairments are especially at increased risk of longer hospitalizations. Mild impairment on the ADL was not generally associated with increased risk of hospital use.

Among the covariates, having been hospitalized during the year prior to the survey and being White were the only covariates related to subsequent hospital use. Previous hospitalization was the only covariate predictor of subsequent stays of six or more days (OR 2.73; CI = 1.96, 3.81).

To directly compare the SPMSQ with the ADL and IADL as predictors of hospitalization, additional logistic regression models were constructed. One set of models included both the SPMSQ and the ADL. Another set included the SPMSQ and the IADL (the ADL and IADL were too highly correlated to have included both of these functions in the same regression models). In these direct comparisons, the SPMSQ proved to be more predictive of future hospital use than the ADL but not as good as the IADL. Five or more dependencies on the IADL was an especially strong predictor of future hospital use even when controlling for moderate to severe impairments on the SPMSQ.

Discussion

The data are consistent with previous findings that cognitive impairment in community dwelling elderly is more common among females, older age groups, and those living with others.[16,17] The prevalence rates of cognitive impairment in studies of community elderly vary from a low of 1.6 percent (severe impairment) to a high of over 15 percent for the mildly impaired. The present study found an overall rate of 16.7 percent. The prevalence of cognitive impairment among the old old (85 years of age and older) was found to be higher (32.8 percent) than that previously reported (25 percent).[16] Two recent studies have also found higher than expected prevalence rates in the old old. Forty seven percent of those 85 years old

TABLE 4—Odds Ratios for Hospital Use by Level of Impairment

		Any Hospital Use		Hospitalized 6 Days or More	
	Percent With Impairment	Odds Ratio	(95% CI)	Odds Ratio	(95% CI)
ADLs					
Mild	26.7	1.30	(.92, 1.82)	1.42	(.97, 2.07)
Moderate/Severe	14.1	1.52	(.99, 2.33)	1.87	(1.19, 2.93)
IADLs					
Mild	45.0	1.52	(.87, 2.68)	2.95	(1.31, 6.64)
Moderate	26.8	2.04	(1.13, 3.69)	3.45	(1.50, 7.96)
Severe	16.8	3.48	(1.84, 6.55)	7.36	(3.12, 17.38)
SPMSQ					
Mild	9.3	1.85	(1.14, 2.98)	1.89	(1.14, 3.16)
Moderate/Severe	7.4	1.76	(1.03, 3.01)	2.10	(1.20, 3.68)

NOTE: Odds Ratios adjusted for age, education, race (White/non-White), and any hospitalizations during year prior to baseline survey.

These data categorize IADL dependencies as follows: 1-2 = mild; 3-4 = moderate; 7.5 = severe.

or older were found to have probable Alzheimer's disease in one study conducted in a geographically defined community in which 80.8% of the age-eligible residents were evaluated.[18] In another study of dementia in a retirement community, Pfeffer[19] found the prevalence rate of dementias to be 56.7 percent in those 85 years old and older. Although comparisons among various studies are difficult due to differing methods of case ascertainment, sampling methods and study populations, these data suggest that the public health impact of cognitive impairment in the elderly may be greater than previously described.

Our data do not support the concept of a racial or ethnic difference in cognitive impairment as implied in earlier suggestions for adjusted scoring of the SPMSQ.[10] The data did show a significant association between educational level and performance on mental status examinations. This association has been documented in other studies as well.[20] For example, Katzman[21] studied a probability sample of 5,055 community dwelling elderly in Shanghai. The Chinese Mini-Mental State (CMMS) was used to evaluate mental status. In comparison to those with formal education, this study found significantly lower scores on the mental status test as well as a different error pattern among the uneducated even after making adjustments for education. The reason for the association between educational level and suggested cognitive impairment has not yet been fully explained.[22]

Although the cause of the cognitive impairment and hospitalization were not determined in this study, the data indicate that cognitive impairment is a risk factor for hospital admissions and longer than average hospital stays even after controlling for age, ADL, IADL, and prior hospitalizations. The longer length of stay may be due to lack of appropriate posthospital placement or to complications of hospitalization.[23,24] Whatever the reasons for increased hospital use, the increasing prevalence of cognitively impaired elderly will result in an enormous financial burden for the patient, family, and government and has significant public health consequences.[25,26] □

Acknowledgments

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The German Health Care System: What Can We Learn from It?

As health costs continue to rise and many Americans find themselves without adequate health care coverage, attention has turned to other countries for answers. In recent years, Canada's government-run "single-payer" system has been the subject of intense debate. In the past year, though, interest has grown in the German "all-payer" health care system. Could it become a model for reforming health care in the United States? The May issue of *The Internist: Health Policy in Practice*, published by the American Society of Internal Medicine, examines what health care lessons can be learned from the German experience.

Uwe E. Reinhardt, PhD, the James Madison Professor of Political Economy at Princeton University, writes that if organized American medicine wishes to get ready for health care by the end of the century, it will need to study the "all-player" system of Western Europe in which "all players in a market area are bound by common fee schedules and all providers are subject to overall budget constraints." Dr. Reinhardt, who is a member of the Physician Payment Review Commission, adds that a move toward an all-payer system in the United States would require the development of a mechanism in which payers, including the federal and state government, could jointly negotiate binding fee schedules with the providers of health care.

German economist J. Matthias Graf von der Schulenburg observes that the annual increase of per capita health care spending in Germany was one of the lowest among industrialized nations between 1980 and 1987 (followed only by Japan, which more or less copied the German health care system). More than 1100 "statutory sickness funds"—self-sustaining, nonprofit institutions generally organized by geographic location, place of employment, or trade or profession—provide care for the German population. Although Germany saw a sudden increase in health care costs in the mid-1970s, Dr. Graf von der Schulenburg examines how costs came under control through negotiations at the different levels (federal, state, and local) of the health care system.

Although Germany's health care system has had a good record in restraining costs and prices, there have been trade-offs in terms of physician supply and demand, says Dale Rublee, PhD, a policy analyst with the American Medical Association. The growth in supply has become the medical profession's greatest long-range concern and Germany's most visible problem, he says. Germany has the most medical students per capita (next to Italy) and has seen rapid growth of medical specialists. The trade-off with an oversupply of physicians has been an overall decrease in physician earnings in the past few years in Germany.

Robert B. Doherty writes in *Capital Currents* that recent studies reporting that a quarter of all health care spending goes to "administrative waste" is overstated and that the data, especially concerning overhead, need to be closely examined. Doherty, ASIM's vice president of governmental affairs and public policy, adds that one of the reasons overhead costs may seem to be higher in the United States than in other countries is that more dollars are invested in quality assurance (such as peer review organizations or continuing medical education requirements).

For a copy of the May issue of *The Internist: Health Policy in Practice*, send $3 to ASIM, 1101 Vermont Avenue, NW, Suite 500, Washington, DC 20005-3457.

Cause-Specific Mortality in a Population-Based Study of Diabetes

ABSTRACT

Background. Mortality from vascular diseases has been reported to be high in diabetic persons.

Methods. To evaluate mortality from these and other specific causes, we examined cause-specific age-sex standardized mortality ratios in a geographically defined population of younger onset (diagnosed before age 30 and taking insulin, n = 1200) and older onset (diagnosed after age 30, n = 1772) diabetic persons followed for 8.5 years. Cause of death was determined from death certificates.

Results. In younger onset persons, age-sex standardized mortality ratios were significantly high ($P < .05$) for all causes of death (7.5) as well as for diabetes (191), all heart disease (9.1), ischemic heart disease (10.1), other heart disease (6.3), nephritis and nephrosis (41.2), accidents (2.9), and all other causes (3.2). In older onset persons, age-sex standardized mortality ratios were significantly high for all causes of death (2.0) as well as for diabetes (16.8), all heart disease (2.3), ischemic heart disease (2.3), other heart disease (2.1), stroke (2.0), and pneumonia and influenza (1.7).

Conclusions. Diabetic persons experience very high mortality, especially from vascular diseases, compared to the general population. (*Am J Public Health.* 1991;81:1158–1162)

Scot E. Moss, MA, Ronald Klein, MD, MPH, and Barbara E. K. Klein, MD, MPH

Introduction

Diabetic persons have been shown to be at greater risk of mortality from cardiovascular disease,[1-18] stroke,[1,2,5,9,11] and renal disease[14,19] compared to nondiabetic persons. Risk estimates for cancer have been variable.[1,3,11,14,15,20] Many of these earlier results are based on clinic populations, include only people with type I or type II diabetes, include males only, or are based on small numbers of cases. Thus, the results of many of these studies may not be applicable to a wider diabetic population.

In the current investigation we examine data from a population-based study, the Wisconsin Epidemiologic Study of Diabetic Retinopathy. Cause-specific mortality from this sample of diabetic persons is compared to the general population.

Methods

The study population has been described in detail in previous reports.[21-23] The study area is composed of 11 counties in southern Wisconsin (1980 population of 839 324). Diabetic persons were identified by a review of the records of 452 of the 457 physicians providing primary care to diabetic persons in the year beginning July 1, 1979. Chart reviews were done on 9841 patients. Of these, 558 were excluded because they were confined to nursing homes, had died prior to July 1, 1979, had incorrect diagnostic codes, had moved, or had gestational diabetes.

A two-part sample of 2990 of the remaining 9283 patients was selected on July 1, 1980, for examination. The first part consisted of all persons diagnosed before age 30 and taking insulin, referred to as younger onset (n = 1210). The second part consisted of a probability sample of persons diagnosed as having diabetes after age 30 confirmed by a random or postprandial serum glucose of at least 11.1 mmol/L or a fasting serum glucose of at least 7.8 mmol/L on at least two occasions, referred to as older onset (n = 1780). The specifics of the sampling strategy and exact proportions appear in an earlier publication, but approximate proportions were 25% for eligible persons with 0 to 4 or 5 to 14 years duration and 100% for persons with 15 or more years of diabetes.[23] Of the older onset group, 824 were taking insulin and 956 were not. Because it was not possible to examine the entire population, these sample sizes were chosen to ensure that an adequate number of cases of diabetic retinopathy would be found to evaluate risk factors for this complication.

Of the original sample, 18 (10 younger onset and 8 older onset) had died or were lost prior to July 1, 1980. These persons were excluded from further analysis. The remaining persons were contacted annually by telephone to determine vital status. In addition, relatives, designated contact persons, and physicians were contacted, and newspaper obituaries were reviewed daily. An annual request was made to the Wisconsin Center for Health Statistics, Section of Vital Statistics, for death certificate information. Wisconsin death rec-

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All authors are with the Department of Ophthalmology, University of Wisconsin Medical School, Madison, Wisconsin.

Requests for reprints should be sent to Dr. Ronald Klein, Department of Ophthalmology, Clinical Science Center, 600 Highland Avenue, Madison, WI 53792.

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ords through 1988 have been searched. Finally, information on persons who moved out of Wisconsin and were suspected of being deceased and persons lost to follow-up was submitted to the National Death Index at the National Center for Health Statistics for matching against national death data through 1987.

The study period was from July 1, 1980, through December 31, 1988. Only deaths that had been confirmed by a death certificate were included in the analysis. Persons who were thought to be deceased but for whom a death certificate could not be located were considered to be alive as of the last contact date they were known to be alive. An analysis was undertaken at this time because nearly half of the older onset group had died, affording adequate deaths to examine a number of underlying causes. While the number of deaths in the younger onset group is considerably smaller and should probably be regarded as preliminary, it is still large enough to investigate the major causes of death.

All medical conditions on the Wisconsin death certificate were coded by trained nosologists in the Wisconsin Division of Health using the International Classification of Diseases, Ninth Revision (ICD-9). The underlying cause of death was selected by the Automated Classification of Medical Entities computer program.[24,25] Out-of-state certificates were coded and processed in the same manner. The cause-specific mortality analysis of the present investigation was based on the underlying cause of death.

Standardized mortality ratios (SMR) were used to compare the mortality experience of the two diabetes groups with that of the general Wisconsin population.[26] The Wisconsin rather than the US population was used as the standard because it is closer in racial and ethnic composition to the diabetic population, which is predominately non-Hispanic White. Age-, sex-, and cause-specific deaths and age- and sex-specific population estimates for the standard population were obtained from published tables.[27,28]

Results

TABLE 1—Vital Status of Younger and Older Onset Diabetic Persons, Wisconsin Epidemiologic Study of Diabetic Retinopathy, 1980–1988

	Younger Onset		Older Onset	
Status	n	%	n	%
Confirmed dead	145	12.1	859	48.5
Suspected dead[a]	5	0.4	17	1.0
Alive at last contact	1039	86.6	883	49.8
Lost to follow-up[b]	11	0.9	13	0.7
Total	1200	100.0	1772	100.0

[a] Treated as alive as of last contact at which known to be living.
[b] No contact for more than 2 years.

Males comprised 52.2% of the younger onset group and 45% of the older onset group. Mean age at diagnosis and standard deviation (SD) in the younger onset group was 14.6 ± 7.6 years. In the older onset group, mean age ± SD at diagnosis was 55.5 ± 12.6 years. Mean age ± SD at the start of the study was 28.4 ± 13.3 and 66.7 ± 11.5 years in the younger and older onset groups, respectively.

The vital status of the younger and older onset groups is presented in Table 1. There have been 145 (12.1%) and 859 (48.5%) confirmed deaths in the younger and older onset groups, respectively. Only 5 (0.4%) younger onset and 17 (1%) older onset deaths have not been confirmed. In addition, only 11 (0.9%) younger onset and 13 (0.7%) older onset persons have been lost to follow-up.

Table 2 presents the distribution of underlying cause of death among diabetic persons for the 12 most common causes of death in Wisconsin. In the younger onset group, diabetes was the leading underlying cause of death, accounting for 44.1% of all deaths. This was followed by heart disease (30.3%) and accidents (6.2%). In the older onset group, heart disease (48.8%) was the leading underlying cause of death, followed by diabetes (15.3%), malignant neoplasms (9.9%), and cerebrovascular disease (9.5%).

Standardized mortality ratios (SMR) by cause of death in the younger onset group are shown in Table 3. Mortality from all causes was 7.5 times more than expected from the general Wisconsin population and was more in excess in females (8.9) than in males (6.8). Among specific underlying causes of death, mortality from diabetes was 191 times greater than that expected. The SMR for heart disease was 9.1, with the excess being greater in females (10.3) than in males (8.7). The difference was more pronounced for ischemic heart disease (SMR of 13.5 in females, 9.1 in males). However, excess mortality for other heart disease was greater in males (7.3) than in females (4.5). There was also a sex difference with respect to mortality from accidents. The SMR for females (5.9) was significantly different from unity while in males it was not (2.0). The only other statistically significant excess mortality was due to nephritis and nephrosis (41.2). There were too few deaths from this cause to separate the sexes. Other SMRs, while substantially greater than unity, were not statistically significant, possibly due to a lack of power because of the small numbers of observed and expected deaths (Table 3).

Table 4 displays SMRs by cause of death in the older onset group. Mortality from all causes was twice the expected, and the excess was slightly greater in females (2.2) than males (1.9). As in the younger onset group, the greatest excess was due to diabetes (16.8). Mortality from heart disease was 2.3 times the expected and affected the sexes equally. Excess mortality due to ischemic heart disease was slightly higher in males while that due to other heart disease was higher in females. Cerebrovascular disease mortality was twice the expected. The only other statistically significant excess mortality was due to pneumonia and influenza (1.7).

Discussion

Many of the studies of cause-specific mortality among diabetic persons compared to nondiabetic persons are of clinic populations,[2,5,14,20] industrial or worker groups,[1,4,9,10,17] or other groups that are not population-based.[11] Clinic populations of diabetic persons may be biased in that they may be more likely to be suffering from complications of diabetes and have a poorer survival experience. Indeed, we have shown that in terms of all-cause mortality, diabetic persons with ocular diseases are at a disadvantage compared to those without.[29] On the other hand, studies based on worker groups may experience bias due to the "healthy worker" effect. Thus, population-based studies or studies based on a representative sample of a population offer the best opportunities for assessing the effect of diabetes on mor-

TABLE 2—Distribution of Cause of Death in Younger and Older Onset Diabetic Persons, Wisconsin Epidemiologic Study of Diabetic Retinopathy, 1980–1988

		Younger Onset		Older Onset	
Cause of death	ICD-9 Codes	n	Cause of Death (%)	n	Cause of Death (%)
Malignant neoplasms	140–208	4	2.8	85	9.9
Diabetes mellitus	250	64	44.1	131	15.3
Heart disease	391–398,402,404–429	44	30.3	419	48.8
Ischemic heart disease	410–414	36	24.8	326	38.0
Other heart disease	391–398,402,404,415–429	8	5.5	93	10.8
Cerebrovascular disease	430–438	3	2.1	82	9.5
Atherosclerosis	440	0	0	10	1.2
Other arterial disease	441–448	0	0	6	0.7
Pneumonia and influenza	480–487	2	1.4	27	3.1
Chronic obstructive pulmonary disease	490–496	1	0.7	12	1.4
Chronic liver disease and cirrhosis	571	2	1.4	7	0.8
Nephritis and nephrosis	580–589	4	2.8	8	0.9
Accidents	800–949	9	6.2	12	1.4
Suicides	950–959	4	2.8	1	0.1
All other causes	—	8	5.5	59	6.9
Total	—	145	100.0	859	100.0

TABLE 3—Observed (O) and Expected (E) Deaths, Age- and Sex-standardized Mortality Ratios (SMR), and 95% Confidence Intervals (CI) by Cause of Death and Sex in Younger Onset Diabetic Persons, Wisconsin Epidemiologic Study of Diabetic Retinopathy, 1980–1988

	Male			Female			Total		
Cause of Death (ICD-9 Code)	O/E	SMR	95% CI	O/E	SMR	95% CI	O/E	SMR	95% CI
Malignant neoplasms (140–208)	—	—	—	—	—	—	4/5.1	0.8	0.2–2.0
Diabetes mellitus (250)	37/0.19	199.6	140.5–275.1	27/0.15	180.3	118.8–262.3	64/0.34	190.9	147.0–243.8
Heart disease (391–398, 402, 404–429)	31/3.6	8.7	5.9–12.3	13/1.3	10.3	5.5–17.7	44/4.8	9.1	6.6–12.2
Ischemic heart disease (410–414)	25/2.8	9.1	5.9–13.4	11/0.81	13.5	6.7–24.2	36/3.6	10.1	7.1–13.9
Other heart disease (391–398, 402, 404, 415–429)	6/0.82	7.3	2.7–15.9	2/0.44	4.5	0.5–16.3	8/1.3	6.3	2.7–12.5
Cerebrovascular disease (430–438)	—	—	—	—	—	—	3/0.74	4.1	0.8–11.8
Atherosclerosis (440)	—	—	—	—	—	—	0/0.05	0	—
Other arterial disease (441–448)	—	—	—	—	—	—	0/0.16	0	—
Pneumonia and influenza (480–487)	—	—	—	—	—	—	2/0.26	7.6	0.9–27.4
Chronic obstructive pulmonary disease (490–496)	—	—	—	—	—	—	1/0.40	2.5	0–13.9
Chronic liver disease and cirrhosis (571)	—	—	—	—	—	—	2/0.42	4.8	0.5–17.2
Nephritis and nephrosis (580–589)	—	—	—	—	—	—	4/0.10	41.2	11.1–105.5
Accidents (800–949)	5/2.4	2.0	0.7–4.8	4/0.68	5.9	1.6–15.2	9/3.1	2.9	1.3–5.5
Suicides (950–959)	—	—	—	—	—	—	4/1.4	2.8	0.8–7.3
All other causes	5/1.6	3.2	1.0–7.4	3/0.9	3.3	0.7–9.6	8/2.5	3.2	1.4–6.4
Total	89/13.1	6.8	5.5–8.4	56/6.3	8.9	6.7–11.5	145/19.4	7.5	6.3–8.8

Note. Sex-specific SMRs are not given where there are fewer than five total deaths.

tality. The present population-based study is notable for its substantial population of both younger and older onset diabetic persons of both sexes.

The population-based studies carried out in the United States have all found an excess mortality from cardiovascular disease.[3,6,7,12,16,18] These studies found SMR and relative risks (RR) in the range of 1.7 to 4.9. Although SMR and RR are not strictly comparable, they may give similar results in practice. The studies cited include both insulin- and noninsulin-dependent diabetic persons and thus are probably most comparable with our older onset group. In our older onset group we found an SMR of 2.3 for cardiovascular disease. This agrees remarkably well with the other population-based studies.[3,6,7,12,16,18] In addition, several of these studies note a differential effect of sex on cardiovascular disease

TABLE 4—Observed (O) and Expected (E) Deaths, Age- and Sex-standardized Mortality Ratios (SMR), and 95% Confidence Intervals (CI) by Cause of Death and Sex in Older Onset Diabetic Persons, Wisconsin Epidemiologic Study of Diabetic Retinopathy, 1980–1986

	Male			Female			Total		
Cause of Death (ICD-9 Code)	O/E	SMR	95% CI	O/E	SMR	95% CI	O/E	SMR	95% CI
Malignant neoplasms (140–208)	38/49.0	0.8	0.5–1.1	47/41.1	1.1	0.8–1.5	85/90.2	0.9	0.8–1.2
Diabetes mellitus (250)	49/3.2	15.4	11.4–20.4	82/4.6	17.6	14.0–21.9	131/7.8	16.8	14.0–19.9
Heart disease (391–398, 402, 404–429)	216/94.9	2.3	2.0–2.6	203/89.4	2.3	2.0/2.6	419/184.3	2.3	2.1–2.5
Ischemic heart disease (410–414)	179/73.2	2.4	2.1–2.8	147/65.8	2.2	1.9–2.6	326/139.0	2.3	2.1–2.6
Other heart disease (391–398, 402, 404, 415–429)	37/21.7	1.7	1.2–2.3	56/23.6	2.4	1.8–3.1	93/45.3	2.1	1.7–2.5
Cerebrovascular disease (430–438)	30/17.1	1.8	1.2–2.5	52/23.4	2.2	1.7–2.9	82/40.6	2.0	1.6–2.5
Atherosclerosis (440)	6/3.3	1.8	0.7–4.0	4/4.3	0.9	0.3–2.4	10/7.6	1.3	0.6–2.4
Other arterial disease (441–448)	3/3.6	0.8	0.2–2.5	3/2.2	1.4	0.3–4.0	6/5.8	1.0	0.4–2.3
Pneumonia and influenza (480–487)	14/8.7	1.6	0.9–2.7	13/7.6	1.7	0.9–2.9	27/16.3	1.7	1.1–2.4
Chronic obstructive pulmonary disease (490–496)	10/9.9	1.0	0.5–1.9	2/4.5	0.4	0–1.6	12/14.4	0.8	0.4–1.5
Chronic liver disease and cirrhosis (571)	5/1.9	2.6	0.8–6.2	2/1.2	1.7	0.2–6.2	7/3.1	2.3	0.9–4.7
Nephritis and nephrosis (580–589)	3/2.3	1.3	0.3–3.9	5/1.9	2.6	0.8–6.1	8/4.2	1.9	0.8–3.8
Accidents (800–949)	3/4.0	0.7	0.2–2.2	9/2.9	3.1	1.4–5.9	12/6.9	1.7	0.9–3.0
Suicides (950–959)	—	—	—	—	—	—	1/2.0	0.5	0–2.8
All other causes	31/21.6	1.4	1.0–2.0	28/24.4	1.1	0.8–1.7	59/46.0	1.3	1.0–1.7
Total	409/221.1	1.9	1.7–2.0	450/206.0	2.2	2.0–2.4	859/429.1	2.0	1.9–2.1

Note: Sex-specific SMRs are not given where there are fewer than five total deaths.

mortality in diabetic persons wherein females are at greater risk relative to nondiabetic persons than males.[6,7,12] However, this finding is not universal.[16,18] We also find no sex differential for either all heart disease or ischemic heart disease in the older onset group, but we do find one in the younger onset group, especially with respect to ischemic heart disease. However, in the Framingham study a trend is noted where the sex differential in cardiovascular disease incidence diminishes with advancing age.[6] If this effect carried over to mortality, it could explain our result.

Several studies have also investigated cerebrovascular disease mortality among diabetic persons.[1,2,5,9,11] Although none of these studies are population based, most of them found varying degrees of excess cerebrovascular disease mortality among diabetic persons. We found a significant twofold excess mortality from cerebrovascular disease in our older onset group. In the younger onset group there is a fourfold excess cerebrovascular disease mortality, but it does not reach statistical significance. However, the SMR is based on only three deaths.

Cancer was investigated in several studies with varying results.[1,3,11,14,15,20] In some studies cancer mortality did not differ from expected,[1,3,11,14] in one it was higher than expected,[15] and in one it was lower than expected in males only.[20] In two of these the results are based on small numbers of deaths,[1,14] and only two are population based.[3,15] We did not find that cancer mortality differed from the expected in either the younger or older onset groups.

Renal disease mortality has been found to be high in younger diabetic persons,[14] but the effect decreases sharply with age.[19] We also found excess mortality due to renal disease in the younger onset group. Mortality due to renal disease in the older onset group, while nearly double that expected, was not statistically significant. However, the real impact of renal disease on mortality in our population may be underestimated because some cases may also be included under diabetes as ICD-9 code 250.4, Diabetes with renal manifestations. If we add these to the other cases of renal disease, then the SMR for nephritis and nephrosis in the younger onset group is 72.1 (95% CI = 28.9–148.5), and in the older onset group it is 2.6 (95% CI = 1.3–4.7).

Two studies reporting on mortality from accidents and suicides combined, showed lower than expected mortality in males and higher than expected mortality in females.[2,14] In the younger onset group, we found higher than expected mortality from accidents in both males and females. In the older onset group, males had lower than expected and females had higher than expected mortality from accidents. In both groups only the result for females was statistically significant. Also, we found higher than expected suicide mortality in the younger onset group and lower than expected in the older onset group, but neither is statistically significant.

The only other significant result is higher than expected mortality due to pneumonia and influenza in the older onset group. We have not found this specific cause of death reported elsewhere. However, because we have a large number of deaths, especially in the older onset group, we are able to examine more specific causes of death than other population-based studies with smaller samples of diabetic persons. ☐

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Hypertension Management in Health Care for the Homeless Clinics: Results from a Survey

Kraig Kinchen, BS, and James D. Wright, PhD

ABSTRACT

Background. With the exception of alcohol abuse, hypertension is the most common chronic physical health problem encountered among homeless persons. The material conditions of homelessness greatly complicate the management of this disorder. Some of the complications and their solutions are discussed here, based on the experiences of health clinics for the homeless in large US cities.

Methods. In 1988, the Stewart B. McKinney Homeless Assistance Act established health care clinics for homeless persons in 108 cities. We surveyed medical directors in these clinics, asking about the management of hypertension in this difficult-to-treat population; 65 responded.

Results. Comparisons between our survey data and those obtained in two recent surveys of clinicians in "normal" clinical practice provide interesting lessons in how medical practice is adapted to respond to the unique needs and problems of the urban homeless. Although therapeutic goals are similar, the means chosen to achieve them often are not.

Conclusions. The treatment of homeless hypertensives illustrates the problems inherent in strict biomedical models of disease and its alleviation. Preferred treatments, course of disease, and success of intervention are powerfully affected by social factors. (*Am J Public Health*. 1991;81:1163–1165)

Introduction

With the exception of alcohol abuse, hypertension is the most common chronic physical health problem encountered among homeless persons, afflicting between 14% and 25% of this population.[1–4] The management of hypertension in the homeless is complicated by material and existential conditions. Proper dietary practice is exceedingly difficult to maintain,[5] rates of alcohol abuse exceed 50%,[4] storage and administration of medicines are difficult, and the cost of medication can be prohibitive. Compliance and maintenance of contact with homeless hypertensives are also difficult, and the asymptomatic stages of the disease produce little or no motivation to seek or maintain treatment.

In 1984, the Robert Wood Johnson Foundation, sensing the widespread unmet medical needs of the nation's homeless, funded health care for the homeless (HCH) clinics in 19 cities. In 1988, the McKinney Homeless Assistance Act extended the system to an additional 89 cities. We surveyed these 108 HCH clinics to learn how medical practice is adapted to the unique needs of homeless hypertensives, using a questionnaire loosely modeled on two recent surveys of hypertension management in "normal" clinical practice.[6,7] One of the clinics served only children and was dropped from the sampling frame; responses were obtained from 65 (60.7%) of the remaining 107 clinics.

Results

Initiation of Therapy

The Joint National Committee for the Detection, Education, and Treatment of High Blood Pressure (JNC)[8] recommends initiation of drug therapy when diastolic blood pressure (DBP) exceeds 94 mm Hg or at somewhat lower pressures if other risk factors are present; nondrug therapies are recommended for DBP between 90 and 94 mm Hg. HCH physicians largely adhere to the JNC recommendations. For an otherwise asymptomatic 45-year-old homeless white male, nondrug therapy would be initiated at an average DBP of 90.5 mm HG or at an average systolic blood pressure (SBP) of 151 mm Hg. Likewise, drug therapy would usually be initiated at a DBP of 97 mm Hg and at an SBP of 161 mm Hg. In both respects, HCH physicians differ little from clinicians in Maryland and New York City. Therapeutic goals are also very similar in both settings, with DBP <90 mm Hg as the target for the overwhelming majority.

The importance of factors other than blood pressure in deciding to initiate therapy was rated on a 1 (not important at all) to 4 (very important) scale. The most important other factor is target organ damage (mean rating [MR] = 3.88). (In a sample of homeless hypertensives being treated in the New Orleans HCH clinic, 20% had concurrent heart problems, compared to 2% of the normotensives.[2]) The co-occurrence of diabetes mellitus (MR = 3.79), elevated blood cholesterol (MR = 3.39), and being a smoker (MR = 3.32) were also rated as important factors in the decision to initiate therapy.

The authors are with the Department of Sociology, Tulane University, New Orleans.

Requests for reprints should be sent to James D. Wright, Department of Sociology, Tulane University, 220 Newcomb Hall, New Orleans, LA 70118-5698.

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TABLE 1—Percentage of Physicians Performing Various Laboratory Tests in the Initial Diagnostic Evaluation of a Recently Diagnosed Hypertensive Man

Laboratory Test	New York City	Maryland	HCH
Aldosterone	5	NR	0
Blood Urea Nitro	95	NR	85
Chest radiograph	100	70	34
CBC	100	67	66
Electrocardiogram	100	89	63
Intravenous pyel	19	11	5
Plasma renin	0	7	2
Serum cholesterol	95	NR	89
Serum creatinine	90	84	92
Serum potassium	100	82	86
Triglycerides	67	NR	55
Uric acid	90	NR	60
Urine culture	24	NR	0
Urinalysis	100	NR	94

Note. HCH = Health Care for the Homeless; NR = not reported.
Sources: New York City[7]; Maryland[8]; HCH, authors' survey.

Laboratory Protocols

Initiation of therapy usually implies a detailed laboratory workup to assess other risk factors and to establish baseline values.[8] Table 1 compares laboratory protocols of HCH physicians with those in Maryland and New York. Urinalysis, serum creatinine, serum cholesterol, and serum potassium tests are the most common laboratory tests ordered in all three settings but all tests except serum creatinine and potassium tests are less likely to be ordered in HCH settings, apparently because most HCH clinics do not have their own laboratories. One respondent noted, "Obtaining initial lab tests in our program requires sending a homeless hypertensive patient to our central health center." Resistance to treatment and absence of a fee source often preclude sending homeless hypertensives off-site for laboratory work. Routine tests are seen as needless complications or unaffordable luxuries.

Nondrug Therapies

We asked HCH physicians how frequently they recommended various nondrug therapies; the Maryland survey asked a somewhat different question that, nonetheless, provides interesting comparisons. In the Maryland study, the most important nondrug therapies, in descending order of importance, were sodium reduction, weight reduction, smoking reduction, increased exercise, and alcohol reduction. Among HCH physicians, alcohol reduction is by far the most important, followed at some distance by sodium reduction, weight reduction, and fat restriction. Alcohol abuse is strongly associated with blood pressure levels in all patient populations[9,10]; the greater importance accorded alcohol reduction by HCH physicians doubtlessly reflects the exceptional rate of alcohol abuse, especially among homeless men.

Dietary control of hypertension is obviously problematic in any population that depends on soup kitchens and bread lines for its sustenance. As one homeless hypertensive in New Orleans put it, "You gotta eat what they put on the plate." Since the prospects of dietary control for homeless patients are dim, they are accorded less importance in HCH clinics than elsewhere. The same can be said of other nondrug therapies often employed in conventional settings. Relaxation therapy is obviously pointless when maintaining the most basic necessities of life is a daily struggle. Smoking reduction may be equally pointless when cigarettes are the only remaining pleasure in life. The barriers to managing homeless hypertensives with nondrug therapies usually mean that antihypertensive drugs are the treatment of first *and* last resort.

Drug Therapies

Most HCH clinicians (45%) consider diuretics the treatment of choice, followed by calcium channel blockers (26%), beta blockers (12%), acetylcholinesterase inhibitors (11%), and "all others" (3%). Cost and availability of medications are more important factors in treating homeless hypertensives because the regimen is often life-long and therefore expensive. Nearly four fifths of HCH physicians cited cost as an important factor. Said one, "Newer and often more effective medicines can't be initiated because the homeless and the agencies that care for them can't afford them. The older medicines, with frequent dosing intervals and side effects, are the only ones we have available." Cost is not usually cited as a critical factor in HCH clinics that provide free medication.

Treatment and Compliance

A recent study of homeless hypertensives in New Orleans found that 59% had their blood pressures under control at least once in the course of treatment and 40% were in control at the last visit[2]; 60% of the domiciled hypertensives in a San Francisco study had normal blood pressure readings at their last visit.[11] Thus, effective treatment of homeless hypertensives is difficult but not entirely hopeless.[1,12,13]

One difficulty in treating the homeless is noncompliance, but this is also a problem for domiciled hypertensives.[14,15] Still, 91% of HCH physicians felt that compliance was more difficult for the homeless. Perceived barriers to compliance are summarized in Table 2. Alcohol abuse and availability of medication require no further comment. In domiciled populations, "The amount of social support a patient receives is strongly correlated with adherence to treatment" (p 355)[14]; among the homeless, estranged and disrupted social networks are common.[16] Further, approximately one homeless person in three is psychiatrically impaired.[17] Homeless patients frequently have trouble keeping appointments because they lack the routines of family and work, and many have no sense at all of day, month, or even year.

Strategies

How do HCH clinicians cope with the many difficulties summarized above? As one noted, with "Perseverance, patience, and creativity." To reduce confusion about medication, many HCH physicians prescribe once-daily dosages; to maintain contact, they schedule more frequent appointments and write prescriptions for shorter periods. Most HCH clinics also employ outreach workers to monitor clients in the streets. At least one HCH clinic has a "no-wait" drop-in station to provide immediate blood pressure readings. Many HCH facilities provide in-clinic medication storage. When possible, medication is dispensed in weather-resistant packages. At least three HCH clinics have experimented with clonidine transdermal patches, which eliminate the storage and daily-dosage problems altogether.[12,13]

TABLE 2—Relative Importance of Nine Factors in Impeding Compliance among Homeless Hypertensives: HCH Physician Assessments

Factor	Rated Importance[a]
Alcohol abuse	3.79
Unavailability of medication	3.76
Lack of social support network	3.42
Psychiatric impairment	3.40
Difficulty keeping appointments	3.34
Resistance to life-style changes	3.24
No place to store medication	3.10
Inadequate understanding of the effects of hypertension	3.08
Side effects of medication	2.63

Note. HCH = Health Care for the Homeless

[a] On a 1–4 scale where 1 = not important and 4 = very important.

One powerful incentive for hypertensives to maintain contact is that other needs can also be addressed at the HCH clinic. Most clinics have social workers, case managers, substance abuse counselors, and other service workers as well as medical personnel. Many of these clinics have evolved essentially into "one-stop" social service centers. Professional nutritionists can help with dietary management; 23% of the HCH clinics had a nutritionist on staff and an additional one third had access to nutritionists via referral.

The biggest barrier to treatment of homeless hypertensives, of course, is that their overall physical and mental well-being is threatened by much more pressing concerns than elevated blood pressure. Given their lack of food, clothing, shelter, and money, there is little concern among the homeless with preventive health measures or health maintenance.

Comment

Homelessness is a social condition with profound consequences for physical well-being, including the effective management of chronic physical disorders such as hypertension. The treatment of homeless hypertensives illustrates the folly inherent in strict biomedical models of disease and its alleviation. The range of treatments available, the course of the disorder, and the success of intervention are all powerfully affected by social factors.

It is an unfortunate commentary that we invest as much in providing health services for the homeless as we do in eradicating the basic causes of homelessness. People cannot be socially, mentally, or physically healthy without first having a stable and secure place to live. Health care clinics help homeless people lead more comfortable, less degrading lives, but they do little to reduce the number of homeless persons. Health care for the homeless, however creative, sensitive, and professionally motivated, cannot and does not address the more basic problems. The ultimate folly would be to become so proficient in dealing with the health consequences of homelessness that we lose sight of the fundamental problem of homelessness. ☐

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Epidemiologic Studies Utilizing Surveys: Accounting for the Sampling Design

Edward L. Korn, PhD, and Barry I. Graubard, MA

ABSTRACT

Background. Since large-scale health surveys usually have complicated sampling schemes, there is often a question as to whether the sampling design must be considered in the analysis of the data. A recent disagreement concerning the analysis of a body iron stores-cancer association found in the first National Health and Nutrition Examination Survey and its follow-up is used to highlight the issues.

Methods. We explain and illustrate the importance of two aspects of the sampling design: clustering and weighting of observations. The body iron stores-cancer data are reanalyzed by utilizing or ignoring various aspects of the sampling design. Simple formulas are given to describe how using the sampling design of a survey in the analysis will affect the conclusions of that analysis.

Results. The different analyses of the body iron stores-cancer data lead to very different conclusions. Application of the simple formulas suggests that utilization of the sample clustering in the analysis is appropriate, but that a standard utilization of the sample weights leads to an uninformative analysis. The recommended analysis incorporates the sampling weights in a nonstandard way and the sample clustering in the standard way.

Conclusions. Which particular aspects of the sampling design to use in the analysis of complex survey data and how to use them depend on certain features of the design. We give some guidelines for when to use the sample clustering and sample weights in the analysis. (*Am J Public Health.* 1991;81:1166–1173)

Introduction

Many biomedical questions involving human populations cannot be answered by experimentation. We cannot ethically randomize patients to certain risk factors, e.g., exposure to asbestos. For factors we can manipulate, we may be unwilling to wait the required number of years or to follow the large number of individuals necessary for a meaningful analysis. Epidemiologic studies offer an alternative to randomized experiments that can avoid the major ethical problems and sometimes allow reduced sample sizes and earlier completion times.

Depending on the risk factors and outcomes, epidemiologic studies may still require large sample sizes. Large multipurpose health surveys can be a readily available source of data. An advantage of this type of data is that many risk factors can be examined. For example, consider the National and Hispanic Health and Nutrition Examination Surveys—NHANES I, II, III and HHANES—conducted in 1971–1975, 1976–1980, 1988–1994, and 1982–1984, respectively. Table 1 gives some examples of cross-sectional associations that have been studied using these surveys. Also, because these surveys were done at different times, one can examine cross-sectionally changes over time in various measured quantities (Table 1). In 1982–1984, a follow-up survey was conducted of NHANES I sampled individuals aged 25 to 74 years. This follow-up survey allows true longitudinal studies to be conducted (see Table 1 for examples). With diminishing financial resources, we believe that these types of surveys will be used increasingly to investigate biomedical and public health hypotheses.

A question that usually arises in the analysis of such survey data is whether the complex sampling design, which involves the clustering and weighting of observations, needs to be accounted for, or if it can be ignored by treating the sampled data as the population of interest. In some[5,7–12,15–20,22,23,25–36] of the above mentioned studies the sampling design was used in the analysis, while in others[1–4,6,13,14,21,24,37–53] it was not. The recent analysis by Stevens et al.[41] that ignored the sampling design in the analysis crystallizes the issue. Among other findings, Stevens et al.[41] found statistically significant differences in the mean total iron-binding capacity and mean transferrin saturation between men who developed cancer and men who did not, controlling for age and smoking status. Yip and Williamson[54] criticized this study because the sampling design was not used, and suggested that if it had been incorporated in the analysis, "it is doubtful that [these differences] would have remained 'statistically significant'." Stevens and colleagues replied, "It is important to take into account the probability sampling method used by NHANES when one is attempting to estimate the level of a variable in the US population at large. To test for differences between case patients and controls within the NHANES cohort, however, the methods we used are appropriate."[55] What is the correct approach to the analyses of these types of studies?

Edward L. Korn is with the Biometric Research Branch and Barry I. Graubard is with the Biometry Branch, National Cancer Institute.

Requests for reprints should be sent to Edward L. Korn, Biometric Research Branch, National Cancer Institute, Executive Plaza North, Room 739, Bethesda, MD 20892.

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TABLE 1—Examples of Studies Using Data from the National Health and Nutrition Examination Surveys

Outcome	References	Risk Factors/Subgroups
Cross-sectional associations		
Blood pressure	1–8	Nutrient intake; alcohol intake; blood lead levels; anthropometric measurements; race, sex, and age
Body weight	9–10	Smoking; alcohol
Anxiety	11	Aspects of diet
Periodontal disease	12	
Tooth decay	13	
General health	14	
Serum cholesterol	15	
Dietary supplements	16	
Blood iron levels	17	
Dietary pattern	18	
Constipation	19	
Glucose tolerance	20	Oral contraceptives; job characteristics
Cardiovascular disease	21–22	
Blood iron	23–24	Race, sex, and age
Blood lead	25	
Serum vitamin A	26	
Serum retinol	27	
Serum α-tocopherol	28	
Dietary intake	29	
Low birth weight	30	
Changes over time measured cross-sectionally		
Blood pressure	31	
Blood lead levels	32	
Anthropometric measurements	33	
Oral contraceptive use	34	
Incidence of childhood asthma	35 36	
Longitudinal studies		
Incidence of cancer	37–44	Alcohol intake; serum cholesterol; dietary fat; anthropometric measurements; blood iron; bowel function; physical activity; depression
Cardiovascular disease mortality	45	Cardiac size
Mortality from injuries	46	Alcohol consumption
Mortality (all causes)	47–50	Diabetes; race; antihypertensive drugs; self-rated health status
Depression	51	Physical activity
Hip fracture	52	Anthropometric measurements
Psychological well-being	53	Age

This article offers a brief tutorial about the sampling design and analysis of large household surveys. We examine whether the sampling design should be used when analyzing cross-sectional or longitudinal data derived from such surveys. We do not consider case-control studies, for which it is generally agreed that the sampling design must be considered in the analysis.[56–58] We consider conditions under which the sampling design can be safely ignored and the inefficiency resulting from utilizing the sampling design when it was unnecessary. The NHANES I and its follow-up are used as illustrations throughout, with special emphasis on the body iron stores-cancer association found by Stevens et al.[41]

Design of Complex Surveys: Cluster Sampling and Sample Weights

In this section the NHANES I design is used to describe the complex sampling design of household surveys. We stress the points of the design relevant to the issues discussed in this paper; other sources[59–61] provide further details. Although simple random samples are the easiest to analyze, they are impractical for large-scale national surveys for two reasons. First, no national registry of names to be sampled may exist. Second, even if a simple random sample of names could be obtained, it would be impractical to travel to, say, 20 000 locations scattered throughout the United States to interview the sampled individuals. Multistage survey designs, on the other hand, enable sampling frames to be developed for a relatively small number of areas and limit the number of locations to which interviewers must travel. For example, in the first stage of NHANES I sampling, the mainland United States was divided into approximately 1900 geographic areas or primary sampling units (PSUs), each one consisting of a standard metropolitan statistical area or, at most, three contiguous counties. These PSUs were grouped into 40 strata from each of which one or two PSUs were sampled; some very large PSUs were sampled with certainty. In each of these sampled PSUs, successive stages of sampling involved census enumeration districts (approximately 300 households), segments (approximately 8 contiguous housing units), and sampled individuals.

At first glance accounting for this type of multistage clustering in the analysis of survey data seems a formidable task. Fortunately, approximate analytic methods have been developed that require only the knowledge of the stratum and PSU of each individual.[62] In practice, computer data files from large surveys contain this information so that it can be readily accessed by statistical computer packages. For NHANES I, the design can be approximated[60] for the statistical analysis as a selection of 70 PSUs, two from each of 35 strata.

For complex surveys, each sampled individual with data has a sample weight associated with his or her data. The sample weight is the number of individuals in the target population that the sampled individual represents. The sample weight may be derived as the product of three components. The first comes from the fact that surveys frequently over-sample certain groups in the population (base weight). For example, in NHANES I, people living in poverty census enumeration districts were sampled as much as eight times more often than people living elsewhere. Additionally, persons aged 65 years or older were sampled twice as often as women ages 20–44, who in turn were sampled twice as often as other adults.

TABLE 2—Adjusted Mean Values[a] for Total Iron-Binding Capacity for Subjects in Whom Cancer Developed and Those Who Remained Cancer-free

	Clustering Used?	Sample Weights Used?	Cancer[b]	No Cancer[c]	Difference		
					Mean±SE	95% CI	P
				Men			
A[d]	No	No	61.30	62.78	−1.48±0.60	−2.65 to −0.31	.013
B	No	Yes	62.65	63.57	−0.92±1.06	−2.99 to 1.15	.38
C[e]	Yes	No	61.30	62.78	−1.48±0.54	−2.58 to −0.38	.010
D[e]	Yes	Yes	62.65	63.57	−0.92±0.86	−2.67 to 0.83	.29
E	Yes	Partly[f]	62.07	63.51	−1.44±0.53	−2.51 to −0.37	.010
				Women			
A[d]	No	No	66.32	66.40	−0.08±0.79	−1.62 to 1.46	.92
B	No	Yes	65.60	66.32	−0.73±1.18	−3.04 to 1.58	.54
C[e]	Yes	No	66.32	66.40	−0.08±0.69	−1.48 to 1.33	.91
D[e]	Yes	Yes	65.60	66.32	−0.73±1.29	−3.34 to 1.88	.58
E	Yes	Partly[f]	66.20	66.35	−0.16±0.60	−1.38 to 1.06	.82

Note. SE = standard error; CI = confidence interval.

[a] Means are adjusted for age and smoking status. For the analysis labeled E, means are also adjusted for race (White vs non-White), family income (<$3000, $3000–$6999, $7000–$9999, $10 000–$14 999, ≥$15 000), poverty vs nonpoverty census enumeration district, and senior status (age ≥65 or <65). For the 3.8% of the sample whose family income was unavailable, income was imputed from education of the head of household in the same manner used in the construction of the sample weights.

[b] For men, N = 242; for women, N = 203.

[c] For men, N = 3116; for women, N = 5165.

[d] Analyses labeled A are from Stevens et al.[41] Adjusted means for cancer and no cancer groups are slightly different here because their analysis implicitly assumed that all four smoking categories were equally likely. Analyses here use observed proportions in each smoking category (analyses A and B) or estimated population proportions in each category (analyses C, D, and E).

[e] Mean estimates for analysis C and D are the same as estimates for analysis A and B, respectively.

[f] Analyses labeled E are unweighted regressions with means adjusted for many of the variables used in defining sample weights (see footnote a).

Finally, in a small number of segments there were more households encountered by the interviewers than expected. These households were subsampled at rates of one half to one quarter.

The second component of the sample weight is an adjustment for nonresponse, including both the inability to locate sampled individuals and their refusal to participate. Although there is no ideal approach to adjust for nonresponse, adjustment using weights is as follows. The population is divided into groups so that the probability of nonresponse is thought to be roughly similar for all individuals within a group. The data from respondents is then weighted upwards depending on the nonresponse rate of their corresponding group. For example, in NHANES I the nonresponse groups were based on five family income groups within each PSU.

The third component of the sample weight is an adjustment so that the sum of the weights for a given sex, race, and age agree with known population figures (poststratification adjustment). Finally, because not all sampled persons in NHANES I had their body iron stores measured, a special set of weights was derived for use in the analysis of these data.

Reanalysis of the Body Iron Stores-Cancer Association

In this section we follow the analyses given in Stevens et al.[41] with the exception of using the sample clustering and/or the sample weights. We demonstrate how the conclusions of the analysis change when these two aspects of the sampling design are incorporated. We restrict attention to the associations described by Stevens et al.[41] of total iron-binding capacity (mol/L) and transferrin saturation (%) with the development of cancer. Briefly, the data to be analyzed are from 3358 men and 5368 women who (1) had their total iron-binding capacity measured in NHANES I in 1971–1975, (2) had their cancer status determined in the follow-up survey in 1981–1984, and (3) were alive and cancer-free for at least 4 years after their biochemical measurements. These sample sizes are slightly different from those of Stevens et al.[41] because the follow-up public use data files have been revised since their analyses.

Although there are many ways to analyze this type of follow-up data, we follow the simplest analysis given by Stevens et al.,[41] which compares the mean biochemical values in those respondents who developed cancer and those who did not, adjusted by linear regression for age (years) and smoking status (current smoker, former smoker, never a smoker, and unknown). Tables 2 and 3 contain the results of five different analyses of these same associations, labeled A through E.

The A analyses repeat the Stevens et al.[41] analyses. We conclude from the A analyses that for both variables there is a small but statistically significant association for men and no association for women.

The B analyses use the sample weights of the survey but not the clustering. A dramatic effect of using the weights is the increase in the standard errors (SEs), the most extreme case being the change from 0.79 to 2.53 for the SE of the difference in adjusted means for the women's transferrin saturation. The correct interpretation for this particular weighted analysis is not that there is no association between transferrin saturation and the development of cancer in women. Instead, it is that because the SE and confidence interval (CI) are so large, the data on the association are uninformative.

The C and D analyses use the sample clustering, without and with the sample weights, respectively. Only the SEs, not the means, are possibly affected by use of the clustering. For these particular analyses, it turns out that the standard errors are not much affected. The results and conclusions, therefore, are very similar to the unclustered A and B analyses, respectively.

We recommend the E analyses for several reasons. They are unweighted analyses that use the sample clustering, but for which the adjusted means have been further adjusted by linear regression for many of the variables used in defining the sample weights (see footnote a of Table 2). The results and conclusions of the E analyses turn out to be very similar to the unweighted unclustered A analyses. This happens to be true even though these analyses do not ignore the sampling weights and clustering.

The computer programs REG[63] and SURREGR[64] were used to perform the analyses A and B–E, respectively. Because 3 of the PSUs have no total iron-binding capacity data available, for the C, D, and E analyses the strata containing these PSUs are pooled with neighboring strata to yield a design approximated by 67 PSUs selected from 32 strata. The analyses A–E give an idea of the different conclusions that are possible when using various aspects of the sample design in the analysis. In the sections below, we give some rules of thumb for choosing the most appropriate analyses based on the sample survey design alone.

Should Sample Clustering Be Used in Analysis?

Sample clustering can affect the variability (SEs) of estimates of associations. First, a simple but extreme example will show this. Consider a simple random sample of 1000 households, where disease and risk factor status are obtained from each member of the household. Suppose that if one member of a household has the risk factor, then all members of the household have it, and if one member of the household has the disease, then all have it. The appropriate sample size for the analysis of this sampled data is 1000, not the number of individuals sampled. Thus, for example, if 4000 individuals had been sampled in the 1000 households, and the sample (household) clustering was ignored in the analysis, then SEs would be too small by a factor of 2 ($= \sqrt{4000/1000}$), because SEs are inversely proportional to sample sizes. Admittedly, this is an extreme example to make the point; in general the effect of clustering depends on both the cluster size and the intraclass correlation.[62]

TABLE 3—Adjusted Mean Values[a] for Transferrin Saturation (%) for Subjects in Whom Cancer Developed and Those Who Remained Cancer-free

	Clustering Used?	Sample Weights Used?	Cancer[b]	No Cancer[c]	Difference		
					Mean±SE	95% CI	*P*
				Men			
A[d]	No	No	33.15	30.73	2.42±0.78	0.89–3.94	.002
B	No	Yes	32.55	30.43	2.12±1.48	−0.78–5.03	.15
C[e]	Yes	No	33.15	30.73	2.42±0.86	0.68–4.16	.008
D[e]	Yes	Yes	32.55	30.43	2.12±1.22	−0.35–4.59	.09
E	Yes	Partly[f]	32.92	30.48	2.44±0.85	0.71–4.17	.007
				Women			
A[d]	No	No	27.98	27.17	0.81±0.79	−0.73–2.36	.30
B	No	Yes	31.37	27.74	3.63±2.53	−1.32–8.58	.15
C[e]	Yes	No	27.98	27.17	0.81±0.76	−0.72–2.35	.29
D[e]	Yes	Yes	31.37	27.74	3.63±2.42	−1.28–8.53	.14
E	Yes	Partly[f]	28.36	27.57	0.78±0.74	−0.72–2.28	.30

Note: SE = standard error; CI = confidence interval.

[a] Means are adjusted for age and smoking status. For the analysis labeled E, means are also adjusted for race (White vs non-White), family income (under \$3000, \$3000–\$6999, \$7000–\$9999, \$10 000–\$14 999, ≥\$15 000), poverty vs nonpoverty census enumeration district, and senior status (age ≥65 or <65). For the 3.8% of the sample whose family income was unavailable, income was imputed from education of the head of household in the same manner used in the construction of the sample weights.

[b] For men, N = 232; for women, N = 197.

[c] For men, N = 3058; for women, N = 5073.

[d] Analyses labeled A are from Stevens et al.[41] Adjusted means for cancer and no cancer groups are slightly different here because their analysis implicitly assumed that all four smoking categories were equally likely. Analyses here use observed proportions in each smoking category (analyses A and B) or estimated population proportions in each category (analyses C, D, and E).

[e] Mean estimates for analysis C and D are the same as the estimates for analysis A and B, respectively.

[f] Analyses labeled E are unweighted regressions with means adjusted for many of the variables used in defining sample weights (see footnote a).

In many realistic situations, the "disease" may be a continuous outcome (e.g., blood pressure) and the risk factor may also be a continuous variable (e.g., dietary calcium intake). Additionally, the risk factor-disease association of interest may be the one in which certain covariates are controlled for, e.g., sex and age. A general sufficient condition for ignoring the clustering in the analysis is the distribution of the outcome for given levels of the risk factor and covariates does not depend on which cluster the individual is in. Given the complicated multistage sampling in a survey such as NHANES I, it may not be obvious whether this condition is satisfied or not.

Since adjusting for the sample clustering will on average increase the SEs, its use in the analysis can be thought of as a conservative procedure. What is lost by using the clustering in the analysis when it was unnecessary? Unnecessary clustering leads to an inefficient analysis, that is, an analysis with CIs wider than necessary and with less statistical power for rejecting the null hypothesis. An easy way to express this inefficiency is in terms of relative sample sizes. For example, an analysis that is 20% inefficient with respect to another analysis means that a sample size of 1000 in the inefficient analysis is equivalent to a sample size of 800 in the more efficient analysis. A key number for estimating the inefficiency of using clustering when unnecessary is the degrees of freedom, d, available for the estimation of SEs. The d is usually taken to be the number of sampled PSUs minus the number of strata. For example, in the NHANES I analysis of body iron stores and cancer, d = 35 = 67 − 32.

The approximate inefficiency in doing an analysis that uses the sample clustering when unnecessary is:[65]

$$\text{Inefficiency} = 1 - (z^{1-\alpha/2}/t_d^{\,1-\alpha/2})^2$$

where $z^{1-\alpha/2}$ and $t_d^{\,1-\alpha/2}$ are the $1 - \alpha/2$ percentiles of a normal distribution and a t-distribution with d degrees of freedom, respectively, and α is the significance level for testing or one minus the confidence level for CIs. For example, for the analyses of NHANES I given in the previous section in which we used 95% CIs for the mean differences, suppose we are also interested in hypothesis testing at the $\alpha = .05$ level. Because d = 35, any standard set of statistical tables gives $z^{0.975} = 1.96$ and $t_{35}^{\,.975} = 2.03$. The inefficiency equals 6.8%, an inefficiency so small that one should use the clustering in the analyses. If the clustering is important, it must be used so that the variability of means is properly estimated; if the clustering is unimportant, little is lost by using it unnecessarily. The similar magnitude of the SEs for analyses A and C in Tables 2 and 3 suggest that, in fact, clustering was not important for these data.

In general, the inefficiency of using the sample clustering when unnecessary becomes greater with decreasing degrees of freedom. For example, the design of HHANES can be approximated[66] by sampling 16 PSUs from 8 strata. For this design, the inefficiency is 28% when the clustering is unnecessary (using d = 8 in the above formula), making the decision of whether or not to use the clustering in the analysis more difficult than with NHANES I. These inefficiency calculations are meaningful whenever a single disease-risk factor association is being examined. For simultaneous analysis of many disease-risk factor associations or other multivariate analyses, the inefficiency of using the clustering when unnecessary can increase.[67]

Should Sampling Weights Be Used in Analysis?

It is well known that the unweighted mean of a sample that is not a simple random sample can lead to a biased estimate

TABLE 4—Percentiles of Sample Weights for Individuals Analyzed

Sex	Minimum	5%	25%	Median	75%	95%	Maximum
Men (n = 3358)	477	876	3614	8170	19 504	31 556	135 824
Women (n = 5368)	611	114[illegible]	3162	6684	11 119	24 090	186 062

of the population mean, whereas a weighted mean provides an approximately unbiased estimate. It is not as obvious that an unweighted analysis of a disease-risk factor association can also lead to the wrong conclusions. A simple numerical example given in the Figure shows this to be true. The expected numbers of sampled individuals are given in the Figure for two sampling schemes for the whole population was well as within two subgroups defined by area of residence. On the left, a study using a simple random sample of 10 000 individuals would show on average the correct result that an association exists between the disease and the risk factor: The probability of having the disease is 5.9% without the risk factor and 12.2% with the risk factor, for a relative risk of approximately 2. On the right, a study of the same population that samples nonurban area residents at nine times the rate of urban area residents would on average sample the same number of individuals from each type of area. Unweighted analysis of data from this probability sample suggests incorrectly that there is no association between the disease and the risk factor. A weighted analysis of the data from this probability sample would on average yield the correct association:

$$2.06 = \frac{(9\times245 + 655)/[(9\times245 + 655) + (9\times4655 + 3445)]}{(9\times15 + 85)/[(9\times15 + 85) + (9\times85 + 815)]}$$

A weighted analysis will provide an approximately unbiased estimate of the population association, whereas an analysis ignoring the weights, in general, will not. A general sufficient condition for when the sampling weights can be ignored, however, is as follows: The distribution of the outcome (disease) for given levels of the risk factor and covariates does not depend on any variables used in the sampling design or used to adjust for nonresponse.

Because it may be difficult to verify this sufficient condition, what is the harm in always using the sample weights in the analysis? As with the use of clustering when unnecessary, the use of sampling weights when unnecessary leads to an inefficient analysis. This inefficiency can be defined as

$$1 - (SE_{unwtd}/SE_{wtd})^2$$

where SE_{unwtd} and SE_{wtd} are the unweighted and weighted SEs of the mean difference. For the linear regression models used in the analyses here, the calculation of the efficiency depends on the distribution of the sample weights and independent variables (see DuMouchel and Duncan[68] for the exact formulas). An approximation is given by the following formula:

$$\text{Inefficiency} = 1 - (w_1 + w_2 + \ldots + w_N)^2/[N(w_1^2 + w_2^2 + \ldots + w_N^2)]$$

where $w_1, w_2, \ldots, w_N$ are the sample weights for the N sampled individuals. The inefficiency increases when the sample weights are more variable, and is zero if all the sample weights are identical.

Table 4 presents selected percentiles of the distribution of sample weights for the individuals analyzed here. The range of the weights is very large, with the largest weight 285 (300) times greater than the smallest weight for the men (women). The inefficiencies of using the sample weights when unnecessary are calculated to be 47% and 48% for men and women, respectively. Considering that the unweighted SEs are not insubstantial compared to clinically meaningful treatment differences, these inefficiencies are too large for us to assert that little is lost by using the weights when unnecessary.

There is an additional consideration when using the sample weights in the analysis. If the sample weights do matter, they can make the analysis even less efficient than described above. For example, consider the analysis of the transferrin saturation in women given in Table 3. Comparing analysis A to analysis B, the SE has increased by a factor of 3.2 = 2.53/0.79, suggesting an inefficiency of 90%. A closer look at the data for this analysis is instructive. Out of the 197 women who developed cancer, the woman with the highest transferrin saturation (68.4%) also had the third largest sample weight (103 042) out of all 5368 women. If this woman's sample weight had had a median value of 6684, then the mean difference in the weighted analysis B would have been 1.35 ± 1.16 instead of 3.63 ± 2.53.

Since the inefficiency of performing a weighted analysis is unacceptably high, what do we recommend? One approach is to use an unweighted analysis but to control for the variables utilized in determining the sample weights. In the present context, the E analyses in Tables 2 and 3 compute the means adjusted for many of these variables. One could argue[69] that because these additional variables (family income, race, etc.) are highly unlikely to be on the causal pathway between body iron stores and the development of cancer, it would be useful to control for these variables even if we had a simple random sample. We view this not as the perfect solution, but as a compromise that avoids the inefficiency of a weighted analysis. As it happens, the results of the E analyses are remarkably similar to the unweighted, unclustered A analyses, but, we can now have more confidence in these results. However, for other data sets and other associations this may not be the case.

Discussion

The impact of sampling design on the analysis of complex survey data should always be considered. Which particular aspects of the design to use in the analysis, however, is a subtler question. Classical survey analysts recommend using both the clustering and the sample weights in the analysis.[70,71] Although this is a philosophically defensible position, we have seen above that it can lead to inferences that make the whole study worthless. On the other hand, the position of Stevens et al.[55] that the sampling design can be ignored when addressing these types of questions is tenuous unless one is willing to make strong assumptions. We have attempted in this article to give simple ways of weighing the costs and benefits of accounting for the sampling design in the analysis. For specific recommendations concerning sample clustering and sample weights, we offer the following rules of thumb.

(1) If the number of sampled PSUs minus the number of strata is greater than or equal to 20 (corresponding to an inef-

SIMPLE RANDOM SAMPLING

(N= 10000 sampled from entire population)

Totals

		Disease -	Disease +	Proportion with Disease
Risk	-	9068	572	5.9%
Factor	+	316	44	12.2%
		Total 10000		

Subgroup 1 (Urban Area)

		Disease -	Disease +	
Risk	-	8379	441	5.0%
Factor	+	153	27	15.0%
		Total 9000		

Subgroup 2 (Non-Urban Area)

		Disease -	Disease +	
Risk	-	689	131	16.0%
Factor	+	163	17	9.4%
		Total 1000		

PROBABILITY SAMPLING

(N= 5000 sampled from each subgroup)

Totals

		Disease -	Disease +	Proportion with Disease
Risk	-	8100	900	10.0%
Factor	+	900	100	10.0%
		Total 10000		

Subgroup 1 (Urban Area)

		Disease -	Disease +	
Risk	-	4655	245	5.0%
Factor	+	85	15	15.0%
		Total 5000		

Subgroup 2 (Non-Urban Area)

		Disease -	Disease +	
Risk	-	3445	655	16.0%
Factor	+	815	85	9.4%
		Total 5000		

FIGURE—unweighted analysis of data from a probability sample can lead to incorrect conclusions

ficiency of roughly 10%), then use the clustering in the analysis. Otherwise, more advanced statistical methods[72] to account for the clustering can be applied to roughly approximate the SEs of the parameters of interest.

(2) Calculate the approximate inefficiency of including the sample weights using the simple formula given in the previous section. If this inefficiency is less than 10%, then use the sample weights in the analysis. If this inefficiency is greater than 10%, then consider the effect this inefficiency will have on the SEs compared to an expected clinically meaningful treatment difference. If the effect is not unacceptably large, use the sample weights in the analysis. Otherwise, use an "unweighted" analysis that controls for the variables relating to the design and nonresponse adjustments in the analysis.

We end with two suggestions for the designers and producers of large-scale health surveys that will be used by other investigators. If our recommendations are followed, one will sometimes need to use variables relating to the design and nonresponse adjustments of the survey. Our first suggestion is that all these variables be documented and included in public use data files.

Our second recommendation concerns the design of the surveys themselves. As we have discussed, a small number of sampled PSUs or individuals with extremely high sample weights can make the use of the sample design in the analysis very inefficient. Given the large number of secondary analyses that are performed using these survey data, we suggest that policymakers provide sufficient resources so that the surveys can be designed for efficient secondary analyses using the design. ☐

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Drug with Potential for Treating the Neuropsychological Effects of HIV Infection to Be Tested

The National Institute of Mental Health (NIMH) today announced the beginning of a clinical study of the effectiveness of Peptide T, a drug of potential value in treating the neuropsychological effects of HIV infection and AIDS.

"This is the first study designed to answer questions about the effectiveness of Peptide T by comparing the compound with a placebo in a controlled study design," said Alan I. Leshner, PhD, NIMH acting director.

The first patient started treatment with Peptide T in mid-March. The study, which is jointly sponsored by NIMH and the National Institute of Allergy and Infectious Diseases, is being conducted at the University of Southern California in Los Angeles.

The human immunodeficiency virus is believed to cause problems of concentration and memory in some persons with AIDS and AIDS-related complex by directly affecting the brain.

Peter N.R. Heseltine, MD, the principal investigator, said "in earlier studies, people with these problems have shown improvement in tests of their mental function when they were treated with Peptide T. The purpose of this new study is to discover if Peptide T is safely able to reduce these HIV-caused problems of concentration and memory and if it is safe and effective when used with the antiretroviral AZT (zidovudine)."

Ellen Stover, PhD, director of the NIMH Office of AIDS Programs, said, "this is the first time that a trial of an AIDS therapeutic will be using improvement in mental functioning as the key outcome. We are working toward making progress in treating HIV infection and are looking forward to the collaboration with USC investigators and the results of this study."

Peptide T was developed by NIMH scientists working at the institute's research facilities on the National Institutes of Health campus in Bethesda, Md. The peptide is an artificially produced protein that blocks the attachment of the AIDS virus to human cells by occupying receptors on the cell surface. Its structure is similar to part of the protein coat of HIV.

In the Phase II study, a minimum of 150 men and women who have been infected with HIV will be enrolled in the double-blind, placebo-controlled trial. (Phase II studies are designed primarily to test the efficacy of new therapies.)

The investigators are making special efforts to enroll women and minorities, who have been underrepresented in some earlier AIDS studies.

Initially, participants will receive either Peptide T or placebo for 6 months, and then during the second 6-month period, all study participants will receive the peptide. During the 1-year study, participants may take any medications approved by the Food and Drug Administration for preventing AIDS-related opportunistic infections, and they may continue treatment with antiviral therapies if they are taking them when they enter the study.

Individuals interested in participating in the study should call Charles Hovis, LAC-USC Medical Center, 213/226-4643, or the PHS-sponsored AIDS Clinical Trials Information Service, 1-800-TRIALS-A.

Breast Cancer Screening among Relatives of Women with Breast Cancer

Karen M. Kaplan, MD, Gene B. Weinberg, DrPH, Alden Small, PhD, and Joy L. Herndon, MS

ABSTRACT

Background. National surveys indicate that only 15% to 30% of all women in the general population 50 years of age or older have an annual mammogram.

Methods. We studied first-degree female relatives of women with breast cancer, who are at elevated risk of disease, to describe the breast cancer screening practices of these relatives and to improve their practices through a program of intensive education. We report here the screening behaviors of 2471 relatives of women with breast cancer.

Results. Self-reported behaviors were as follows: 49% performed monthly breast self-examination and 70% had annual breast examinations by a medical professional. Of 983 women 50 years of age or older, 49% had had a mammogram, but only 14% have a mammogram annually. Of women 50 years of age or older who had never had a mammogram, 92% reported never having had one recommended by a medical professional.

Conclusions. Our findings indicate that screening behaviors in relatives of breast cancer patients are not substantially different from those of women in the general population. Enhanced efforts both to educate medical professionals and to encourage women to demand screening mammography are necessary to reduce breast cancer mortality. (*Am J Public Health.* 1991;81:1174–1179)

Introduction

Nearly one in nine women in the United States will have breast cancer in her lifetime. An estimated 175 000 new cases will be diagnosed and an estimated 44 500 deaths from this disease will be recorded during 1991. Breast cancer ranks second only to lung cancer as a cause of cancer mortality in women; yet despite biomedical advances in breast cancer therapy, mortality has remained fairly constant and has even increased slightly in recent years.[1]

Routine breast examination by a physician and breast self-examination (BSE) both have been shown to be more effective than accidental discovery in detecting the disease at an early stage.[2,3] Moreover, detection of early lesions using a combination of professional breast examination and mammography for women ≥50 years of age, and more recently for women 40 to 49 years of age, has been shown to improve survival from breast cancer.[4]

Despite the potential reduction in morbidity and mortality using available breast cancer screening methods, screening measures are underutilized. Recent population-based surveys indicate that 50% of women ≥50 years of age perform monthly BSE, 50% have annual professional breast examinations, and no more than 15% to 30% have annual mammograms.[5-8]

A total of 8% of women 40 years of age or older in the United States have at least one first-degree female relative (mother, daughter, or sister) who has had breast cancer.[9] Nevertheless, little is known about the screening behaviors of these relatives,[10-12] women whose risk of disease is estimated to be twofold to fourfold higher than that for women in the general population.[13] Krischer and colleagues studied the breast cancer screening behaviors of 138 first-degree and 70 second-degree relatives of women with breast cancer and found low prevalence rates of breast cancer screening overall, rates comparable to those of women in the general population.[10] To further investigate this group of high-risk women, the Pennsylvania Cancer Program of the Pennsylvania Department of Health supported six pilot projects from 1984 to 1987. The projects were designed to identify relatives of women with breast cancer, to describe the breast cancer screening behaviors of these relatives, and to improve their behaviors through a program of intensive education.

We report here the screening behaviors of these relatives in five of the pilot projects. We hypothesized that health care providers would more effectively screen women who were known to be at elevated risk of breast cancer and that the prevalence of screening behaviors among this population would be higher than that among women in the general population.

Karen M. Kaplan and Joy L. Herndon were with the Epidemiology Program Office, Centers for Disease Control, Atlanta, Ga. Gene B. Weinberg and Alden Small are with the Pennsylvania Department of Health, Harrisburg, Pa. Dr. Kaplan is now with the Department of Pediatrics, The Pennsylvania State University College of Medicine, Hershey, Pa.

Requests for reprints should be sent to Karen M. Kaplan, MD, Division of Field Services, Epidemiology Program Office, Centers for Disease Control, Atlanta, GA 30333.

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TABLE 1—Prevalence of Breast Self-Examination (BSE), Professional Breast Examination (BE), and Mammography (MAM) by Selected Characteristics, Relative Breast Cancer Screening Project, Pennsylvania, 1984–1987

Characteristic[a]	Total	Ever BSE (%)	Monthly BSE (%)	Ever BE (%)	Annual BE (%)	Ever MAM (%)	Periodic MAM (%)[b]
Overall	2471	83	49	96	70	37	15[c]
Age (years)							
<40	1059	84	49	96	77	42[d]	NA[e]
40–49	424	86	54	97	70	53	17
>50	963	82	47	96	61	49	14
Race							
White	2362	82	47	96	61	48	15
Non-White	106	79	41	97	74	56	11
Residence							
Urban	1091	84	55	97	76	42	19
Mixed	1268	83	46	95	65	33	13
Rural	112	89	24	96	61	31	4
Relation							
Daughter	1231	85	50	96	74	26	17
Sister	1159	82	49	96	65	47	15
Mother	81	78	38	98	68	54	11
Education (years)							
<12	441	79	44	93	53	39	10
12	1119	84	48	96	69	50	15
>12	858	81	52	98	78	59	23
Income($)							
<20,000	843	79	46	95	61	41	11
20,000–40,000	1001	86	51	96	74	55	17
>40,000	414	85	53	100	80	73	27

[a] Data are not available for all behaviors stratified by all characteristics.
[b] Biennial mammograms in women 40–49 and annual mammograms in women ≥50.
[c] Women ≥40 only.
[d] Women 35–39 only.
[e] NA = not applicable.

Methods

All women diagnosed with breast cancer from 1980 to 1986 were identified from cancer registries in eight Pennsylvania hospitals and from the Pennsylvania Cancer Registry in the south-central region of the state. Only women known to be alive at the time of the study were included. The physician who completed the cancer registry data form was contacted for permission to contact the patient; women whose physicians granted permission received letters describing the study and inviting their participation. Women with breast cancer were asked to identify and give permission to contact their mothers, sisters, and daughters, ≥18 years of age, for enrollment in the study.

Data were collected either by telephone or in face-to-face interviews using a structured questionnaire that included demographic characteristics and information about three breast cancer screening behaviors: performance of BSE, breast examinations by a medical professional, and mammography. All information was self-reported; no attempt was made to verify the reported frequencies of screening behaviors by review of medical records.

Because the majority of participating relatives lived in the area of the hospital and because no specific information on residence was collected, the catchment area of the hospital was used as a surrogate for characterizing residence.

The Statistical Analysis System (SAS Institute, Cary, NC) was used for data analysis. Categorical data were compared using χ^2 and the χ^2 test for trend. In reporting stratified data on screening behaviors, only participants with complete risk factor and outcome information were included. To compute the odds ratio for each of the three screening behaviors while simultaneously controlling for the potentially confounding effects of other characteristics, we used unconditional logistic regression (SAS CATMOD with the CLOGITS option).

Results

Demographics

A total of 4012 women with histologically confirmed breast cancer were identified. At the time of the study, 3229 were still alive and were well enough to participate. The physicians of 2975 (92%) of these gave consent for their patients to be contacted. A total of 2050 women (64%) agreed to participate, and 1862 (58%) were interviewed.

Women with breast cancer identified 3860 relatives, of whom 2865 (74%) were invited to participate. A total of 2540 relatives (66%) agreed to participate, and 2539 were interviewed. Data on 68 of these women (2% of the total group of relatives) who reported that they also had had breast cancer were excluded from the analysis. The analysis thus consists of 2471 women, 64% of the relatives of participating women with breast cancer.

Characteristics of participating relatives are summarized in Table 1. The majority were <50 years of age, White, had completed at least 12 years of education, and had a household income >$20,000. A total of 44% of participants were from urban areas, 51% were from areas of mixed type (urban and rural), and 5% were from rural areas. Information about relatives who declined to participate was not available.

Screening Behaviors

A total of 83% of women have done BSE, and 96% have had at least one breast examination by a medical professional (Table 1). A total of 42% of women 35 to 39 years of age, 53% of women 40 to 49 years of age, and 49% of women ≥50 years of age have had at least one mammogram. A total of 28% of women ≥50 years of age reported having had a mammogram within the previous 12 months.

In order to more accurately distinguish routine screening behaviors from breast examinations or mammography done because of breast abnormalities, we also examined periodic performance of breast cancer screening. Table 1 shows the proportions of women reporting monthly BSE, annual breast examinations from a professional, periodic mammography (biennial mammograms for women 40 to 49 years of age and annual mammograms for women ≥50 years of age) for women overall and stratified by age, race, residence, relationship, education, and income. Younger women, women from urban areas, daughters of women with breast cancer, more highly educated women, and women of higher income more often performed monthly BSE, had annual breast examinations from a professional, and had periodic mammography. Although the numbers of non-White

TABLE 2—Mammography and Breast Examination by a Professional in Women ≥50, Relative Breast Cancer Screening Project, Pennsylvania, 1984–1987

	Mammogram within 12 months		
	Yes	No	Total
Professional breast exam within 12 mo.	No. (%)	No. (%)	No. (%)
Yes	263 (41)	375 (59)	638 (67)
No	6 (2)	315 (98)	321 (33)
Total	269	690	959 (100)

Note: N = 959

TABLE 3—Medical Specialist Doing Breast Examination and Proportion of Their Patients ≥50 Years Having Mammograms, Relative Breast Cancer Screening Project, Pennsylvania, 1984–1987

	Any Mammogram			Annual Mammogram		
Medical specialist	Yes	Total patients	Percentage	Yes	Total patients	Percentage
OB/gyn	194	323	60	66	321	21
Internist	38	67	57	16	67	24
General practitioner	138	356	39	27	356	8
Others[a]	99	178	56	30	178	17

[a]Includes physicians of unknown specialty, surgeons, and oncologists.

women in the study population were small, White women performed BSE and had periodic mammography more frequently. Non-White women more frequently had annual professional breast examinations.

Medical Care Encounters

Of 700 women >40 years of age who had never had a mammogram, 89% reported that a mammogram had never been recommended to them by a medical professional. Stratified by age, 83% of women ages 40 to 49 and 92% of women ≥50 years of age reported never having had a mammogram recommended.

The 75 women who had never had a mammogram, but to whom a mammogram had been recommended, were asked why they had not had one. A total of 19% responded "no particular reason," 17% cited high cost as the primary reason, and 13% said that they planned to have a mammogram.

Of 959 women ≥50 years of age for whom information was available, 67% had had a professional breast examination within the previous 12 months. Of these women, 41% had had a mammogram within that period of time compared with 2% of women who had not had a breast examination within 12 months ($\chi^2 = 163.58$; $P < .001$) (Table 2). Nevertheless, 59% of women who had had a breast examination by a professional within the previous 12 months had not had a mammogram within that period of time.

In the same group of 959 women ≥50, we noted the type of medical specialist usually seen for a breast examination stratified by mammography use and annual mammography use (Table 3). A total of 60% of women who usually have a breast examination by an obstetrician/gynecologist had had a mammogram, compared with 57% seen by internists and 39% who usually have breast examinations by a general practitioner (χ^2 for trend = 34.31; $P < .001$). Moreover, only 8% of women usually seen by a general practitioner for breast examinations have annual mammograms, compared with 21% seen by obstetrician/gynecologists and 24% seen by internists ($\chi^2 = 28.24$; $P < .001$). Some women were seen for breast examinations by surgeons, oncologists, or physicians whose specialties were unknown or unreported. These medical practitioners make up the "other" category. The extent to which the practices of physicians in this category are comparable is not known. The differences by medical specialty persisted when the data were stratified by age groups, education groups, and income groups. When we stratified the data by residence, we found the same differences in mammography use based on the type of physician seen for a breast examination among residents of both urban and mixed (urban and rural) settings. There were too few observations for women living in rural areas to make meaningful comparisons among the three types of residences, however.

Multivariate Analyses

We examined BSE, breast examination by a professional, and annual mammography using unconditional logistic regression. The importance of age, income, education, urban setting, family relation, and type of medical professional seen for a breast examination were all analyzed in connection with the three outcome behaviors mentioned above (Table 4). Controlling for the other variables in the model, monthly BSE was significantly associated with urban residence. For the outcome of annual breast examination by a professional, the variables of most importance when controlling for other factors were younger age, higher income, urban residence, and type of medical professional seen for a breast examination. (Women usually seen by obstetrician/gynecologists were more likely to have annual examinations than women seen by general practitioners.) When these same independent variables were analyzed with annual mammograms as the outcome variable, older age, higher income, urban residence, and type of medical professional seen for a breast examination were the variables of most importance. Women usually seen by obstetrician/gynecologists, internists, or physicians in the "other" category were more likely to have annual mammograms than women usually seen for breast examinations by general practitioners.

Adherence to American Cancer Society Screening Guidelines

We examined the behaviors of women in the two age groups 40 to 49 and ≥50 to determine what proportions adhere to American Cancer Society (ACS) guidelines: monthly BSE; annual breast examinations by a professional; and mammography, biennially for women 40 to 49 and annually for women ≥50 years of age.[14] Seventeen percent of women ages 40 through 49 have biennial mammograms, and 14% of women ≥50 have annual mammograms (Table 1). A total of

TABLE 4—Prevalence of Annual Mammogram (MAM), Annual Professional Breast Examination (BE), and Monthly Breast Self Examination (BSE) by Selected Characteristics, Logistic Regression Model, Relative Breast Cancer Screening Project, Pennsylvania, 1984–1987

	Annual MAM		Annual BE		Monthly BSE	
Variable	OR	95% CI	OR	95% CI	OR	95% CI
Relation (base = daughters)						
Sisters	1.743	1.110–2.735	1.276	0.943–1.725	0.906	0.703–1.168
Mothers	1.130	0.452–2.826	1.519	0.798–2.891	0.516	0.289–0.921
Age (years) (base = <40)						
40–49	3.975	2.352–6.718	0.706	0.507–0.982	1.446	1.096–1.910
50 and older	5.212	3.011–9.021	0.494	0.353–0.691	1.156	0.866–1.543
Residence (base = urban)						
Mixed	0.661	0.475–0.920	0.697	0.556–0.875	0.695	0.575–0.842
Rural	0.141	0.018–1.087	0.693	0.417–1.151	0.320	0.193–0.531
Education (base = <12						
12	0.940	0.593–1.488	1.202	0.902–1.600	0.907	0.686–1.200
>12	1.200	0.709–2.030	1.400	0.985–1.991	0.989	0.719–1.361
Income (base = <$20,000)						
$20,000–$40,000	1.794	1.214–2.653	1.394	1.098–1.770	1.202	0.973–1.486
Over $40,000	2.711	1.679–4.375	1.799	1.277–2.533	1.147	0.869–1.512
Physician specialty (base = general practice)						
Ob/gyn	1.774	1.145–2.748	1.608	1.240–2.086	1.165	0.923–1.472
Internal medicine	2.784	1.417–5.471	1.783	0.977–3.253	0.856	0.514–1.426
Other	2.751	1.666–4.541	1.009	0.727–1.400	1.113	0.820–1.510
Total n	2003		2024		2068	

Note: OR = odds ratio; CI = confidence interval.

TABLE 5—Prevalence of Breast Self-Examination (BSE), Professional Breast Examination (BE), and Mammography (MAM), Pennsylvania, 1984–1987, vs Studies of Women in the General Population

Behavior	Percentage Pennsylvania	Percentage General Population
Monthly BSE	49	47–50[5,9]
Annual BE	70	50–55[5–7]
MAM[a]	50	40–67[7–9,16]
Annual MAM[b]	14	15–30[6–8]
MAM ≤1 yr[b]	28	20–41[7,15,17,21]

[a] Women ≥40 years of age.
[b] Women ≥50 years of age.

41% of women 40 to 49 and 32% of women ≥50 adhere to the recommended screening for two of the three practices, monthly BSE, and annual professional breast examinations. Only 11% and 8%, respectively, adhere to all three screening guidelines for their age group.

We compared our findings with those of other surveys of breast cancer screening among women in the general population (Table 5). The prevalence of monthly BSE and mammography use reported in our study is similar to that of women in the general population.[5–8,15–19] The prevalence of annual professional breast examination, 70%, is higher than generally reported.[5–7,9] The proportion of women in the general population who practice a combination of these three screening methods is not known.

Discussion

Our findings indicate that BSE and mammography use in relatives of women with breast cancer, who are at elevated disease risk, do not differ substantially from the behaviors of women in the general population. Despite the higher rates of professional breast examinations reported for this population, barriers to more widespread use of screening mammography exist. Our data do not support the hypothesis that physicians screen selectively and comprehensively those women known to be at elevated risk of breast cancer. Nearly 90% of women who had never had a mammogram reported that a medical professional had never recommended one.

Older women, women of lower socioeconomic status, and women living in rural areas were less likely to report adherence to current ACS breast cancer screening guidelines. Similar risk factors for inadequate preventive health practices in general and for breast cancer screening in particular have been found in national surveys.[9,19–21]

Not surprisingly, we found that women who reported having had a recent professional breast examination more often reported having had a recent mammogram. Yet the majority of women ≥50 seen for a breast examination within the previous year who would have been eligible for screening mammography were not screened, most often because their physicians had not recommended it.

We found differences in compliance with guidelines for mammography use among medical specialties: women who usually had breast examinations by internists or obstetrician/gynecologists were more likely to have mammograms than women who usually had breast examinations by general practitioners. Reasons for the apparent differences in practices among medical specialists are not clear, but such differences may in part reflect type or recency of medical training or participation in continuing medical education. It is also not known whether women in the study population who have breast examinations by general practitioners were less likely to have health insurance that covers screening mammography than were women seen by internists or obstetrician/gynecologists. Bergner and colleagues[22] found that recent medical school graduates and medical school faculty members were more likely to perform a series of cancer screening procedures among asymptomatic patients over the age of 50 for whom they provided primary care. In their survey, obstetrician/gynecologists followed by oncologists most often performed breast examinations and screening mammograms. Information on years since graduation from medical school or faculty status was not available for the physicians of participants in our study.

The failure of physicians to follow established guidelines for breast cancer screening in asymptomatic women was documented in a 1989 Ohio survey of more than 1000 practicing primary care physicians.[23] The survey found that although 98% advise their patients to do BSE, and 78% followed published ACS guidelines for periodic breast physical examinations, only 37% followed guidelines for screening mammograms in asymptomatic women.[23] A similar survey conducted in 1984 found 80% and 11% of physicians adhering to ACS guidelines for breast physical examinations and screening mammography, respectively.[24]

The National Cancer Institute (NCI) has developed the following objective for breast cancer screening for the year 2000: 80% of women 50 to 70 years of age will have both annual breast examinations by a medical professional and annual mammograms.[25] The authors of *Healthy People 2000* suggest that the NCI breast cancer screening goals be incorporated into the Nation's Year 2000 Health Objectives.[26] To achieve these goals, enhanced educational efforts are needed, particularly in promoting mammography. These efforts should target both primary health care providers and women ≥50 years of age. Medical school, residency training, and continuing medical education programs should include instruction on the current approaches to breast cancer screening. The 1989 survey of practicing primary care physicians suggested that misinformation about radiation risks from mammography may deter some physicians from ordering screening mammograms for their patients.[23] In addition, women should be encouraged to demand screening mammography, an approach that has shown utility in stimulating physicians to provide other preventive services.[27–29]

Several limitations of the study design should be noted. First, participation rates were low for women with breast cancer (58%) and their relatives (64%); yet there was no information by which to compare characteristics of participants and nonparticipants. We believe that nonparticipants are unlikely to have better screening behaviors than participants.

Second, in order for women with breast cancer to provide access to their relatives, only those women still alive at the time of the study were eligible breast cancer patients. Whether relatives of women who have died from breast cancer differ from other relatives in their screening behaviors is not known. Further research is needed to better understand factors such as stage of disease at diagnosis, extent of disease, or type of therapy, all of which may affect preventive behaviors in asymptomatic relatives.

Third, all information in the study was self-reported and thus subject to recall bias on the part of the respondents. No attempt was made to verify the reported frequency of professional breast examinations or mammography use by review of medical records. It is therefore possible that some groups of women are more likely to recall breast cancer screening practices better than others.

Fourth, the population studied was primarily White and of relatively high socioeconomic status; the findings may be neither comparable to findings of other population surveys of breast cancer screening behaviors nor generalizable to all populations. Additional studies should focus on the breast cancer screening practices of non-White women and women of lower socioeconomic status who are at elevated risk of breast cancer.

Despite these limitations, our data confirm the findings of previous surveys of women's breast cancer screening behaviors: screening remains underutilized.[5–9,15–18] Whether any of these surveys, including our own, actually measure women's behaviors as opposed to those of their health care providers is not clear. In either case, intensified efforts are needed to enhance the early detection of breast cancer in order to reduce mortality from this disease. ☐

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New Insights into Sudden Infant Death Syndrome

Results from a study published in the May 1 *American Journal of Epidemiology* suggest that certain labor and delivery events may increase the risk of sudden infant death syndrome (SIDS).

SIDS is the leading cause of mortality in the United States for infants 1 to 12 months old, accounting for one to three deaths per 1000 live births. SIDS has been called "a diagnosis of exclusion" because its victims can only be identified through an autopsy that rules out all other explainable causes of death.

This study "is one of the first analytic studies to focus on the intrapartum period and risk factors for SIDS," says lead author Germaine M. Buck, PhD, clinical assistant professor, Department of Social and Preventive Medicine, School of Medicine and Biomedical Sciences, Buffalo, NY.

Buck and her team found a sevenfold increased risk for SIDS associated with vaginal breech delivery and more than twice the risk for SIDS when mothers were in labor 16 hours or longer. Their study population was selected from a 1974 cohort of live births from Upstate New York that included 132 948 resident mothers. Data were abstracted from hospital delivery and vital records for 148 autopsied cases and 355 matched controls.

"The majority of breech SIDS infants were single footling deliveries (a rarer type of breech presentation with the baby emerging one foot first). The more common form of breech delivery called 'frank' presentation, with the baby emerging buttocks first, was not associated with an increased risk for SIDS," Buck says.

"We suspect that the single footling breech presentation may be an indicator for an earlier problem in the fetus' development, while a breech delivery, per se, is not the direct cause of the syndrome," she says. This speculation concurs with previous studies that have suggested SIDS victims may be compromised before birth. As Buck explains, "SIDS babies may have experienced a prior central nervous system insult in utero since they often show signs of subtle growth retardation, including reduced birth weight and length." These effects could be explained by altered blood flow or oxygen supply to the developing fetus, Buck says, but not by labor or delivery events.

"The association between long labor and SIDS must be viewed conservatively," Buck adds. "There is a lot of subjectivity in the woman's reporting of how long she has been in labor when she is first seen in the hospital."

According to Buck, the incidence of SIDS has not diminished as prenatal care has improved in the last several decades. "SIDS is a chief contributor to this nation's high infant mortality rate, and further effort to understand this syndrome should be a public health mandate," she says.

Factors Influencing the Formation of New Human Coronary Lesions: Age, Blood Pressure, and Blood Cholesterol

Wendy J. Mack, PhD, and David H. Blankenhorn, MD

ABSTRACT

Background. The Cholesterol Lowering Atherosclerosis Study, a controlled angiographic trial, has reported that new native coronary artery lesions are significantly reduced by aggressive blood lipid lowering therapy with colestipol plus niacin. To study factors relevant to primary atherosclerosis prevention, we have conducted an epidemiologic analysis of new native coronary lesion formation in placebo-treated patients.

Methods. Univariate and multivariate logistic regression procedures were used to examine age at entry into the study, number of years since bypass, body weight, diastolic and systolic blood pressure, plasma total cholesterol, triglycerides, HDL-cholesterol, LDL-cholesterol, non-HDL-cholesterol, and apolipoproteins A-I,B, and C-III.

Results. Significant univariate protective factors were older age ($P<.006$), reduction of total plasma cholesterol ($P<.040$), and systolic ($P<.024$) and diastolic ($P<.022$) blood pressure. Significant multivariate protective factors were older age ($P<.005$) and reduction in systolic blood pressure ($P<.021$). Blood pressure effects were not associated with use of specific antihypertensive agents.

Conclusions. These data provide additional support for the control of hypertension and reduction of blood cholesterol level for primary and secondary ischemic heart disease prevention. They also indicate the existence of a population at high risk for early coronary lesion formation and the need for improved means to identify such individuals prior to the onset of clinical manifestations of ischemic heart disease. (*Am J Public Health.* 1991;81:1180–1184.)

Introduction

CLAS (Cholesterol Lowering Atherosclerosis Study) was a randomized, placebo-controlled angiographic trial testing the hypothesis that blood lipid lowering with colestipol plus niacin would slow or reverse the atherosclerotic process. CLAS was a secondary prevention trial involving men, ages 40 to 59 years, who had had previous aortocoronary bypass surgery. CLAS angiograms were evaluated for vessel closure, changes in degree of stenosis, and formation of new lesions. We have previously reported significant treatment benefits in overall coronary status of both native vessels and bypass grafts, including reduction of formation of new lesions.[1] Following publication of CLAS results, angiographic trials testing nifedipine[2] and dipyridamole plus aspirin[3] have also reported reduced formation of new coronary lesions, which suggests that pathways early in atherogenesis may be beneficially modified by several different types of therapy. Studies of atherogenesis in animal models and cultured cell systems suggest a variety of ways in which factors initiating lesion formation might differ from those that cause progression of established lesions. The relationship of risk factors to prevalence of human coronary lesions[4] and change in established coronary lesions[5–20] has been extensively studied by angiography, but there are no risk analyses for new coronary lesion formation. To explore risk factor effects on human new coronary lesion formation we conducted an analysis of risk for disease with nondisease in placebo-treated CLAS subjects by comparing subjects who developed new coronary lesions with those who did not. The influence of diet on new lesion formation has been reported previously.[21] We now address the association of clinical variables with the development of new lesions.

Methods

Study Population

Subjects were nonsmoking males, ages 40 to 59 years, with progressive symptomatic atherosclerosis who had had previous coronary bypass surgery. Screening procedures, criteria for exclusion, standard visit protocols, and basic intervention goals were reported earlier.[22] One hundred eighty-eight subjects were randomized into two treatment groups of 94 each; 162 completed the study, 82 of whom were placebo-treated. Average age at entry of all placebo subjects was 54.5 ± 0.5 (SEM) and average blood pressure was 123 mm Hg ± 1.5 (SEM)/80 ± 1.0 (SEM). Thirty percent never smoked and 70% had quit smoking more than 6 months before trial entry. Twenty-four-hour urine specimens were examined for nicotine and cotinine every 6 months to monitor nonsmoking status. Dietary goals for the placebo group were as follows: total fat calories to provide 26% of energy, less than 5% of energy from saturated fatty acids, 10% of energy from polyunsaturated fatty acids, and less than 250 mg cholesterol per day. One-on-one diet counseling was conducted at each clinic

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The authors are with the Atherosclerosis Research Institute, Departments of Preventive Medicine and Medicine, University of Southern California School of Medicine, Los Angeles.

Requests for reprints should be sent to David H. Blankenhorn, MD, RMR 102, 2025 Zonal Avenue, Los Angeles, CA 90033.

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visit with a review of fat and cholesterol intake in computerized 7-day food records from the preceding week. Levels of physical activity were moderate as judged from questionnaires and estimates of energy expenditure from records of caloric intake for the 7 days preceding each clinic visit.

Data Collection

Angiographic. Baseline coronary angiography was performed immediately before randomization. A repeat angiogram was performed at 2 years by the same angiographer, who was blinded to the treatment assignment. Coronary angiogram evaluation procedures have been described in detail previously.[1,22] All lesions greater than 20% diameter stenosis were identified. Film pairs showing identical coronary artery views were mounted side by side on two Vanguard projectors and viewed simultaneously by two expert angiographers who did not know the subject's demographic and clinical characteristics, treatment assignment, or temporal order of the angiograms. The angiographers were not shown ventriculograms and other angiograms. Working independently, the two angiographers identified all visible lesions in one film (assigned at random), recorded this information, and then shared it to reach a consensus. In this manner a total of 1831 lesions were recognized and evaluated, 82% of which were identified independently by both panelists. For the 1501 lesions identified independently by both panelists, agreement within ±10% stenosis occurred in 76%. After reaching a consensus on the first film, panelists examined the second film alongside the first and compared all lesions first separately and then together. After all film pairs were evaluated, the code for film order was broken and lesions recorded in a later angiogram, but not in a first, were tabulated.

Body Weight, Blood Pressure, Plasma Lipids, and Medication History. Body weight, blood pressure, and plasma lipids were measured at every visit.[22] Blood lipids and lipoprotein cholesterol levels were determined by Lipid Research Clinics methodology standardized against reference materials supplied by the Centers for Disease Control.[23] LDL-cholesterol levels were calculated with the equation of Friedewald et al.[24] Baseline levels were the means of the first three screening visits. Postrandomization visits were scheduled every month for the first 6 months and bimonthly thereafter. On-trial values were the average of all postrandomization visits weighted by the scheduled interval between treatment visits. Apolipoprotein determinations were performed by the Oklahoma Foundation for Medical Research.[22] All medications, including over-the-counter drugs, were recorded at semiannual visits (0, 6, 12, 18, and 24 months after randomization).

At study entry 36% of the 82 CLAS placebo subjects reported use of antihypertensive medications and 93% of this group remained on these agents during the trial. The most common drugs and dosage were propranolol 40 mg/day (33% of subjects), hydrochlorothiazide 50 mg/day (13%), and dyazide 50 mg/day (10%). Eight percent of CLAS placebo subjects reported one or more on-trial uses of either nifedipine, diltiazem, or verapamil. Eighty percent of subjects took either aspirin, Ascriptin, or acetaminophen on one or more occasions during the trial.

Statistical Analyses. Univariate logistic regression procedures were used to determine which independent variables were predictive of the development of new lesions. Potentially predictive independent variables included age at entry into the study, number of years since bypass, body weight, diastolic and systolic blood pressure, plasma total cholesterol, triglycerides, HDL-cholesterol, LDL-cholesterol, non-HDL-cholesterol, and apolipoproteins A-I, B, and C-III. Age and years since bypass were analyzed by values at baseline. All other variables were analyzed by values at baseline, as well as mean on-trial, to explore the temporal relationship of these variables to lesion formation. In addition, the difference (baseline minus mean on-trial) was analyzed to assess the relationship between change in these variables and new lesion formation. We chose to compare baseline to mean on-trial values of each independent variable for two reasons. First, given the relatively small sample size of this study and the fact that each subject contributed a potential of 15 on-trial data points for each variable, it was not feasible to assess the significance of each variable at each time point. Second, given the short time period of this study (2 years) relative to the several decades of life, over which atherosclerotic disease develops, we felt that assessing the association of overall on-trial changes relative to prerandomization measures provided a more meaningful analysis. For each independent variable, univariate odds ratios (OR) and associated 95% confidence intervals (CI) were computed[25]; all *P* values for trends in relative risk are based on two-tailed tests. Stepwise logistic regression was used as a final step to arrive at a multivariate model of statistically independent predictors of new lesion formation.

Results

Of the 82 subjects in the placebo group, 18 had developed new lesions in the 2-year period between angiograms, as previously reported.[1] Thirteen among the 18 subjects with new lesions had a single new lesion, 3 had two new lesions, and 1 each had three and four new lesions. Among the 26 new lesions, 15 were in ungrafted patent arteries, 1 was proximal to complete occlusion in an ungrafted artery, 6 were proximal to open grafts, 2 were distal to open grafts, and 2 were distal to closed grafts. Nine of the new lesions were between 20% and 29% diameter stenosis and 6 were between 30% and 39%. Two lesions were 40% diameter stenosis, 4 were 50%, and 1 each were 60%, 70%, 75%, 80%, and 90% stenosis.

With subjects classified by presence or absence of new lesions, mean values of major clinical variables are presented in Table 1. Descriptively, compared to subjects who did not develop new lesions, subjects who developed new native lesions were usually younger at study entry, had recently undergone bypass surgery, had lost less weight on-trial, and had increased both systolic and diastolic blood pressure while on-trial. Those who developed new lesions also showed minor changes in total cholesterol and LDL-cholesterol (less than 1% decrease in both), a 2% decrease in non-HDL-cholesterol, a 5% increase in HDL-cholesterol, and a 10% decrease in triglycerides. In comparison, CLAS placebo subjects who did not develop new lesions demonstrated larger decreases in total, LDL, and non-HDL-cholesterol (6%, 7%, and 7%, respectively), no change in HDL-cholesterol, and an 8% decrease in triglycerides. Subjects developing new native lesions also had a lower average baseline value of apolipoprotein B (average 111.1 mg/dL compared to 124.1 mg/dL in subjects not developing new lesions) and demonstrated an 8% increase in this variable while on-trial. Subjects who did not develop new lesions showed no average on-trial change in apolipoprotein B. Table 1 also presents the univariate ORs and 95% CIs for each variable. All variables have been divided by their sample standard deviation (SD) so that the resulting ORs represent the relative risk for a 1 SD increase in that variable.

TABLE 1—Independent Variable Means by Lesion Group

Variables	New Lesions Mean ± SD (n = 18)	No New Lesions Mean ± SD (n = 64)	SD[a]	OR[b]	95% CI
Age at study entry	51.7 ± 5.0	55.2 ± 15.7	4.4	0.5	0.3–0.8
Years since bypass	2.4 ± 1.5	2.9 ± 2.3	2.3	0.7	0.4–1.3
Body weight (lbs)					
Baseline	178.3 ± 16.7	177.7 ± 23.3	21.9	1.0	0.6–1.7
On-trial	178.1 ± 16.3	176.8 ± 25.2	23.4	1.1	0.6–1.8
Difference	0.2 ± 6.1	0.9 ± 5.8	5.9	0.9	0.5–1.5
Diastolic BP (mm Hg)					
Baseline	80.8 ± 9.3	80.3 ± 8.3	8.5	1.1	0.6–1.8
On-trial	83.4 ± 7.3	80.1 ± 7.1	7.3	1.6	0.9–2.9
Difference	−2.6 ± 4.5	0.2 ± 4.2	4.4	0.5	0.3–0.9
Systolic BP (mm Hg)					
Baseline	118.8 ± 11.2	124.7 ± 12.0	12.0	0.6	0.3–1.1
On-trial	121.4 ± 10.2	123.0 ± 11.5	11.2	0.9	0.5–1.5
Difference	−2.7 ± 7.1	1.6 ± 6.3	6.9	0.5	0.3–0.9
Plasma total cholesterol (mmol/L)					
Baseline	6.3 ± 1.0	6.3 ± 0.9	0.9	1.0	0.6–1.6
On-trial	6.2 ± 1.2	5.0 ± 0.8	0.9	1.3	0.8–2.1
Difference	0.1 ± 0.5	0.3 ± 0.5	0.5	0.5	0.3–1.0
LDL-cholesterol (mmol/L)					
Baseline	4.3 ± 0.8	4.4 ± 0.8	0.8	0.9	0.5–1.5
On-trial	4.2 ± 1.0	4.1 ± 0.7	0.8	1.2	0.7–2.1
Difference	0.0 ± 0.5	0.3 ± 0.5	0.5	0.6	0.3–1.0
HDL-cholesterol (mmol/L)					
Baseline	1.1 ± 0.2	1.1 ± 0.3	0.2	0.8	0.5–1.3
On-trial	1.1 ± 0.2	1.1 ± 0.2	0.2	1.0	0.6–1.6
Difference	−0.1 ± 0.1	0.0 ± 0.1	0.1	0.7	0.5–1.1
Non-HDL-cholesterol (mmol/L)					
Baseline	5.2 ± 1.0	5.1 ± 0.9	0.9	1.0	0.6–1.7
On-trial	5.1 ± 1.3	4.8 ± 0.8	0.9	1.3	0.8–2.1
Difference	0.1 ± 0.5	0.3 ± 0.5	0.5	0.6	0.3–1.0
Triglycerides (mmol/L)					
Baseline	2.0 ± 1.0	1.7 ± 1.0	1.0	1.3	0.8–2.1
On-trial	1.8 ± 0.9	1.5 ± 0.9	0.9	1.3	0.8–2.1
Difference	0.2 ± 0.3	0.1 ± 0.4	0.4	1.1	0.7–1.9
Apo-A-I (mg/dL)					
Baseline	114.9 ± 15.4	115.6 ± 19.0	18.2	1.0	0.6–1.6
On-trial	123.8 ± 11.6	121.9 ± 15.4	15.7	1.1	0.7–1.9
Difference	−8.9 ± 14.9	−6.3 ± 15.8	15.6	0.9	0.5–1.4
Apo-B (mg/dL)					
Baseline	111.1 ± 25.7	124.1 ± 26.7	26.9	0.6	0.3–1.0
On-trial	119.8 ± 24.2	124.1 ± 22.2	22.6	0.8	0.5–1.4
Difference	−8.8 ± 21.1	0.0 ± 21.0	21.2	0.6	0.4–1.1
Apo-C-III (mg/dL)					
Baseline	5.6 ± 2.8	4.6 ± 2.1	2.3	1.5	0.9–2.5
On-trial	5.7 ± 3.2	4.7 ± 2.5	2.6	1.4	0.8–2.2
Difference	0.0 ± 1.7	−0.1 ± 1.4	1.4	1.1	0.6–1.8

Note. Difference = baseline minus mean ontrial value; CI = confidence interval.
[a]SD = Standard deviation for entire sample (N = 82).
[b]Odds ratio = relative risk per SD change of independent variable (relative to no change).

Independent variables showing univariately significant trends in relative risk for new lesion formation are displayed in more detail in Table 2. To show the trend in relative risk, each variable was categorized into four levels using the sample quartiles, and ORs were estimated for each level relative to the first quartile.

Age at entry into the study was protective for new lesions; thus, older men were less likely to develop new lesions. Dividing the total group of 82 men at the median age (54.5 years), the risk of new lesion development for men above the median was 0.2 times that of men below or at the median age (95% CI = 0.1–0.7). Conversely, the risk for new lesion formation in men younger than 55 years at study entry was 4.8 times that of older men. Inclusion of other significant independent clinical and diet variables[21] did not alter the significance of the influence of age on lesion development.

Changes in both diastolic and systolic blood pressure and total plasma cholesterol from baseline to mean on-trial values were also significant univariate predictors of new lesions (Table 2). Decreases in blood pressure and total plasma cholesterol from baseline to on-trial conferred lower risk for development of new atherosclerotic lesions. Other independent variates, including HDL-, LDL-, and non-HDL-cholesterols, apolipoproteins A-I, B, and C-III, triglycerides, body weight, and years since bypass, were not significantly associated with new lesion formation. Finally, variables indicating whether or not a subject had used platelet-active medications, calcium channel blockers, or other antihypertensive medications were not significantly related to new lesion formation.

To investigate the differential effects of lipids by age, interactions between a dichotomized age variable (dichotomized at the median age at study entry) and plasma lipids measured at baseline and on-trial were included in the logistic regression procedure. Two significant interactions emerged in these analyses. The first interaction was an age-on-trial HDL-cholesterol interaction ($\chi^2 = 3.96$; df = 1; $P = .046$), with the parameter estimate indicating a protective effect for HDL-cholesterol confined to subjects older than 55 years. The interaction between age and on-trial triglycerides was also significant ($\chi^2 = 7.00$; df = 1; $P = .01$) with an increased risk for lesion development due to increased triglyceride level occurring in the older age group. In both cases, the main effect of age remained significant after inclusion of these interactions.

To determine which of the univariately significant variables were independently predictive of new native lesions, a stepwise logistic regression analysis was performed (Table 3). Age at study entry was the first independent variable to enter the model; the risk for development of new native lesions was decreased by a factor of 0.4 for every 4.4 years (ISD) of age. After adjusting for age, the change in systolic blood pressure from baseline to on-trial was also significantly predictive of new lesion formation; each 7 mm Hg decrease in systolic blood pressure was associated with the risk of new lesions being decreased by a factor of 0.5. No other variables were significantly related to lesion development after inclusion of these two variables.

Of the 18 subjects who had developed new lesions, 11 had new lesions in native arteries with related grafts, while 7 had new lesions in native arteries without a related graft. These two subgroups were separately compared to the 64 subjects who did not develop new lesions. Parameter estimates for the final model (age at entry into study and change in systolic blood pressure) did not change from those displayed in Table 3, suggesting that these effects are the same in native arteries with and without related bypass grafts.

Discussion

Our analyses are focused on the appearance of new lesions visible by angiography because these indicate an increase in the extent of raised intrusions into the coronary lumen, the most significant single anatomic predictor of ischemic heart disease (IHD) at autopsy.[25] We are concerned primarily with relationships within the placebo group because new lesions were more common here and because these relationships are more relevant to primary IHD prevention than those among drug-treated subjects. Reduction of new lesions in secondary prevention has merit, but reduction of new lesions in primary prevention could benefit many more individuals and might lead to population-wide control of IHD. We have previously reported reduction of new lesions in secondary prevention in CLAS through aggressive treatment with colestipol plus niacin.[1] This treatment reduced average total plasma cholesterol from 6.36 to 4.65 mmol/L and reduced the relative risk of new lesions by a factor of 0.4 ($P = .03$). We have also reported an analysis of diet records of CLAS placebo subjects relevant to primary prevention that indicated that reduction of new lesions might be accomplished by reducing dietary fat.[21] We now report an additional analysis of risk factors relevant to primary prevention that have significant effects on new lesion formation independent of diet. In CLAS placebo subjects, risk for new lesions was predicted by age, on-trial reduction of total plasma cholesterol, and on-trial reduction of both systolic and diastolic blood pressure. When these four univariate predictors competed in stepwise logistic regression, age and change in systolic blood pressure were found to be significant. The effect of age is significant in analyses that pool drug subjects with placebo.

TABLE 2—Univariately Significant Predictors of New Lesion Formation[a]

	OR	95% CI	*P* trend
Age at entry to study, y			
40 to 52	1.0		.006
>52 to 54.5	0.3	(0.1, 1.1)	
>54.5 to 58	0.1	(0.02,0.6)	
>58	0.1	(0.03,0.8)	
Change in total blood cholesterol (mmol/L)			
≤0	1.0		.040
>0 to 0.2	0.8	(0.2, 2.9)	
>0.2 to 0.6	0.6	(0.2, 2.5)	
>0.6	0.2	(0.1, 1.4)	
Change in diastolic blood pressure (mmHg)			
≤−4	1.0		.022
>−4 to −1	0.4	(0.1, 1.5)	
>−1 to 3	0.3	(0.1, 1.5)	
>3	0.2	(0.04,1.1)	
Change in systolic blood pressure (mmHg)			
≤−3	1.0		.024
>−3 to 1	0.6	(0.2, 2.3)	
>1 to 5	0.4	(0.1, 1.8)	
>5	0.2	(0.04,1.3)	

Note. OR = odds ratio; CI = confidence interval.

[a] Change variables = baseline minus mean on-trial.

TABLE 3—Stepwise Logistic Regression

Variable	OR[a]	95% CI[b]	*P*
Age at study entry	0.4	0.2–0.8	.005
Change in systolic blood pressure	0.5	0.3–0.9	.021

[a] Odds ratio interpreted as change in risk per standard deviation of age or standard deviation change in systolic blood pressure.

[b] CI = confidence interval.

The multivariate stepwise model associated with on-trial change in systolic blood pressure and age indicates that reduction of systolic blood pressure by 10 mm Hg reduced the risk of the new lesion development by 60% in men of median age 54.5 years. It is important to note that the principal antihypertensive agents used in CLAS placebo subjects were propranolol, hydrochlorothiazide, and dyazide, not slow calcium channel antagonists. When variables indicating whether or not a subject had taken calcium channel blockers or other antihypertensive medications were introduced into logistic regression, they were not significantly related to new lesion formation. We conclude that reduction of blood pressure level per se, rather than the agent used to produce the reduction, was the significant influence in CLAS placebo subjects. Thus, modest changes in blood pressure perceptibly altered atherogenesis at the stage of raised lesion development. This finding is compatible with the large body of epidemiologic evidence for long-term effects of blood pressure on population IHD rates.[27,28] Trials of antihypertensive therapy such as the Multiple Risk Factor Intervention Trial[29] may have failed to show reduced IHD rates because benefits from reduced new lesion formation were masked by catastrophic events in existing coronary lesions.

In summary, placebo group data from CLAS indicate that formation of new intrusive coronary lesions is significantly influenced by age and change in blood cholesterol and blood pressure level. The effects of blood pressure and blood cholesterol are compatible with findings from IHD epidemiology and confirm that it is appropriate to reduce these risk factors for primary IHD prevention. Age was observed to have a negative effect, with more young subjects showing new lesion formation. This indicates the need for improved methods to identify young individuals who form new coronary lesions at accelerated rates. □

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A Randomized Trial of an Interim Methadone Maintenance Clinic

ABSTRACT

Background. Interim methadone maintenance has been proposed as a method of providing clinically effective services to heroin addicts waiting for treatment in standard comprehensive methadone maintenance programs.

Methods. A clinic that provided initial medical evaluation, methadone medication, and AIDS education, but did not include formal drug abuse counseling or other social support services was established in New York City. A sample of 301 volunteer subjects recruited from the waiting list for treatment in the Beth Israel methadone program were randomly assigned to immediate entry into the interim clinic or a control group.

Results. There were no differences in initial levels of illicit drug use across the experimental and control groups. One-month urinalysis follow-up data showed a significant reduction in heroin use in the experimental group (from 63% positive at intake to 29% positive) with no change in the control group (62% to 60% positive). No significant change was observed in cocaine urinalyses (approximately 70% positive for both groups at intake and follow-up). A higher percentage of the experimental group were in treatment at 16-month follow-up (72% vs 56%).

Conclusions. Limited services interim methadone maintenance can reduce heroin use among persons awaiting entry into comprehensive treatment and increase the percentage entering treatment. (*Am J Public Health*. 1991;81:1185–1191)

Stanley R. Yancovitz, MD, Don C. Des Jarlais, PhD, Nina Peskoe Peyser, MBA, Edmund Drew, MD, JD, Patricia Friedmann, MS, Harold L. Trigg, MD, and J. Waymond Robinson, MD*

Introduction

Expansion of drug abuse treatment programs has been one of the more frequently recommended means for controlling the spread of the human immunodeficiency virus (HIV) and AIDS among intravenous drug users (IVDUs). In the United States, the Presidential Commission on the HIV Epidemic[1] and the National Academy of Sciences[2,3] have both advocated "treatment on demand" (providing immediate voluntary entry into treatment) for IVDUs. The National Institute on Drug Abuse has determined that providing treatment for IVDUs will be its primary strategy for reducing the spread of HIV among drug injectors.[4]

The Food and Drug Administration (FDA) has proposed changing the regulations governing methadone maintenance treatment in the United States to permit interim methadone treatment. The term "interim" refers to the provision of limited services to patients awaiting treatment positions in comprehensive methadone programs. Interim clinics would provide intake physical examinations, education about AIDS, and methadone medication to prevent narcotic withdrawal symptoms and to block the euphoric effects of heroin. The methadone would be dispensed daily by a nurse so that there would be frequent contact between the patient and the medical staff and minimal opportunity for diversion of the medication. Interim clinics would not be required to provide ongoing drug abuse counseling, vocational rehabilitation, or the other social services that are incorporated in current federal regulations governing methadone maintenance treatment and that are integral components of most drug-free programs.

According to their proponents, interim clinics would not replace regular methadone maintenance programs, but would provide limited services to heroin addicts who would otherwise be on waiting lists and receiving no drug abuse treatment. The limited services are expected at least to reduce heroin use and the AIDS risks associated with heroin use. Opponents argue that interim methadone maintenance would not address the multiple social and vocational needs of most heroin addicts, and that implementation of limited service treatment might impede expansion of more comprehensive treatment programs, or, in a time of limited public resources for drug abuse treatment, even replace more comprehensive treatment.[1] (For an extended discussion of the proposed FDA regulations that would have permitted interim methadone maintenance treatment see the transcript of hearings conducted by the Select Committee on Narcotics Abuse and Control of the U.S. House of Representatives, March 23, 1990.)

The intensity of debate notwithstanding, there are relatively few data on the effectiveness of interim methadone clinics. Hong Kong and The Netherlands both have some methadone maintenance pro-

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All authors are with the Beth Israel Medical Center, New York, NY.

*Harold L. Trigg has died since this study was completed.

Requests for reprints should be sent to Stanley R. Yancovitz, MD, Beth Israel Medical Center, New York, NY 10003.

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grams that provide few supportive services beyond basic dispensing of the medication. These programs are generally considered effective within those societies,[5,6] but the effectiveness of interim methadone maintenance might be different in the United States, due to cultural differences with respect to drug use and the higher percentage of polydrug use among US heroin addicts. We report here on a random assignment evaluation of an interim methadone clinic in New York City. Subjects receiving interim clinic services were compared to waiting list controls. Change in heroin use was the primary outcome measure utilized; changes in cocaine use and entry into conventional treatment were also examined.

Methods

Interim Clinic Services

The interim clinic was opened in February 1987. It was staffed by a one-quarter-time physician, one full-time registered nurse, one clinic coordinator, one outreach worker, and one full-time research interviewer. The services provided within the clinic included a standard physical examination upon admission and methadone administered by a nurse 5 days per week. Saturday medication and a single take-home dose were provided at another site in the same building. AIDS education, with free distribution of condoms, was also provided at the interim clinic. Urine samples were collected biweekly for toxicologic analyses, which included thin layer chromatography with confirmation by gas chromatography/mass spectrometry performed by Damon Clinical Laboratories of New York.

Results of the urinalysis were known to the treatment staff, but not incorporated into any treatment plan or discussed in formal counseling sessions, as they generally would be in a comprehensive methadone treatment program. Minimal counseling was available on an ad hoc, informal basis, and other supportive service needs could be addressed only by referral to community agencies. The physician could identify medical problems but, except for the most routine of interventions, had to refer patients elsewhere for treatment.

The initial dose of methadone was set by the examining physician based on the examination and drug use history and was typically 20 to 30 mg/day. This initial dose was then increased according to an escalation schedule preset by the physician, generally by 10 mg every third day until a maintenance dosage of approximately 80 mg/day was reached. The nurse was authorized to modify the rate of dose escalation or to stop the escalation process at a stabilization dose other than the scheduled maximum. Dosage adjustment decisions generally reflected subjective perceptions of the patients and instances of missed or vomited medication.

Subjects and Data Collection

Subjects were recruited from the waiting lists of the Beth Israel methadone maintenance program. Separate waiting lists are maintained for the 23 Beth Israel clinics, and there is substantial variation in the time between application for treatment, placement on a waiting list, and admission to one of the regular clinics. At the time of the interim clinic study, the average time spent on a waiting list was approximately 3 months. Participation in the study would be from the time of enrollment until an opening occurred at the clinic to which the subject had originally applied. At that time the subject would be enrolled in the regular clinic and cease participation in the interim clinic study. Participation in the study would neither decrease nor increase the time spent waiting for an opening in a regular clinic.

A written informed consent for participation in the study was obtained from each subject. The project was approved by Beth Israel's Institutional Review Board and the FDA (Investigational New Drug number 28 232). A questionnaire covering demographics, drug use history, AIDS risk behavior, and knowledge of AIDS was administered by a trained interviewer. The subjects were informed that medical and drug use information collected at the interim clinic would not be revealed to the comprehensive clinic to which they would ultimately transfer. Subjects were paid $35 for the completion of intake data collection. A follow-up questionnaire, focusing on drug use since the preceding interview, was administered every 2 weeks.

Residual serum from the blood sample collected as part of the intake examination was stored at −70 °C for later HIV testing. Since the object of the study was to examine the effects of a methadone treatment program that did not include formal counseling, the actual HIV counseling and testing were postponed until completion of the evaluation study. (Relationships between serostatus and drug use behavior will be presented elsewhere.)

After completion of the initial data collection, subjects were assigned randomly to experimental treatment or control conditions. The random assignment was done by administrative staff at a different location. The intake interviewers frequently asked to have individuals placed in the methadone treatment group based on data obtained in the intake interview, but, consistent with the protocol, these requests were not honored. Subjects in the control condition were paid $20 for follow-up interviews. Data were analyzed using the PRODAS system of statistical programs.[7]

Change in Protocol

For the first 3 months after the study inception, there were three experimental conditions: the experimental treatment group immediately received interim clinic treatment as described above and a biweekly follow-up interview; the frequent contact control group received the biweekly follow-up interview, free condoms, and had biweekly urine samples collected; and the minimal contact control group were not reinterviewed until the end of their participation in the study.

Subject recruitment under these conditions was initially rapid, but then slowed dramatically. Discussions with recruited subjects and potential subjects applying for methadone treatment at Beth Israel indicated that the major difficulty was the perception of a low probability (one chance in three) of receiving methadone treatment in the interim clinic study.

In August 1987, the protocol was revised to increase subject recruitment. The minimal contact control group was eliminated from the study, and the time spent in the frequent contact control group was limited to 1 month, after which control subjects were switched into the experimental group and received methadone medication. After this modification, there was no difficulty in recruiting new subjects for random assignment into experimental and control conditions and the 150-person limit on patients receiving methadone in the interim clinic was soon reached. Data from the discontinued minimal contact control group are not included in this report, but are available from the senior author (S.R.Y.).

Data Collection Limitations

One research staff member was designated to conduct baseline and biweekly research interviews for the experimental and control subjects. The outreach worker, who was bilingual (Spanish/English), occasionally assisted in interviewing. It is thus unlikely that the subjects maintained a clear separation between clinic treatment staff and research staff, which may have contributed to underreporting of drug use as presented below.

TABLE 1—Selected Demographic and Drug Use Characteristics for Subjects at Entry Into Study (N = 301)

Characteristics	Experimental Treatment Group (n = 149) N	(%)	Frequent Contact Group (n = 152) N	(%)	χ^2 or t value	*P* Value
Sex						
Male	121	(81)	118	(78)		
Female	28	(19)	34	(22)	0.59	.44
Race						
White	15	(10)	16	(11)		
Black	44	(30)	61	(40)		
Hispanic	89	(60)	75	(49)	3.93	.14
Currently employed	22	(15)	21	(14)	0.06	.85
Living with sexual partner	59	(40)	62	(41)	0.04	.83
Prior drug treatment	115	(78)	122	(81)	0.43	.51
Prior MMTP treatment[a]	89	(60)	81	(53)	1.27	.26
HIV status						
Positive	65	(50)	93	(64)		
Negative	66	(50)	52	(36)	5.93	<.05
Average age at baseline (yr)	34.8		34.4		t=−0.51	.61
Average age began IV drug use (yr)	18.6		19.4		t=1.15	.25
Average age began IV drug use on regular basis (yr)	19.5		20.2		t=1.02	.31

[a] MMTP = methadone maintenance treatment program.

Given the number of subjects and the severe limitation of clinic space and personnel, some urine collections and interviews were not conducted at the required intervals, especially in the first few months of the study when new intakes and biweekly follow-up interviews for both experimental and control subjects were needed. As the study progressed and the control participants reached their 30th day and were transferred to the experimental treatment group, the total number of study participants stabilized at the maximum of 150 patients receiving methadone, and the rates of missed urines and interviews diminished considerably.

Results

Table 1 shows demographic characteristics and drug use histories for the 301 subjects who were assigned to the experimental and frequent (biweekly) contact control conditions. Ninety-eight percent reported that they had injected drugs at some time in their lives; 92% reported injecting in the month prior to entry into the study, with a mean of over three injections per day. There were no significant differences between groups in demographics or drug use histories. More than half of those who were tested for HIV were seropositive, with a higher percentage seropositive in the control group.

Of the 301 subjects initially enrolled, no follow-up data were obtained from 41 subjects (20 from the experimental group, 21 from the control group). Of these 41 subjects, 32 were lost to contact, 8 were admitted to traditional drug treatment, and 1 was incarcerated before follow-up data could be collected.

Since control subjects were shifted into the experimental treatment after 1 month, primary outcome comparisons between the two groups must be restricted to behavior within that time period. Complete baseline and 1-month follow-up data are available for 169 subjects initially assigned to experimental and frequent contact control groups. The 1-month period was operationally defined as between 15 and 44 days after intake. For subjects with multiple follow-up interviews and urine samples during their participation in the study, the interview and urine sample collected nearest 30 days was used, with the restriction that data from the control subjects reflected only time spent in the control condition. Table 2 presents demographic and drug use history data for these subjects. They do not differ significantly on any of these variables from the total groups assigned to experimental treatment and frequent contact control conditions.

Table 3 presents baseline self-reported drug use and urinalysis results at enrollment for the 169 subjects with complete baseline and follow-up data. There is a general agreement between the self-reports and the urinalysis results. None of the differences between the two groups are statistically significant by chi-square tests.

Table 4 presents self-reported drug use and urinalysis results at the 1-month follow-up. There is clearly substantial underreporting among members of the experimental treatment group. Factors associated with discrepancies between self-reported drug use and urinalysis results will be examined in a separate paper. The drug use analyses presented in this report will be restricted to urinalysis results only. Comparisons of the urinalysis results across experimental treatment and control groups at 1 month using chi-square tests show less heroin use ($\chi^2 = 15.35$, $P < .001$) and more methadone use ($\chi^2 = 52.86$, $P < .001$) in the experimental group. The very small difference in cocaine use was not significant ($\chi^2 = .09$, $P = .7$) between the two groups.

Comparisons from intake to the 1-month time period within each group showed significantly reduced heroin use (63% to 29%, McNemar $\chi^2 = 16.45$, $P < .001$) and significantly increased methadone use in the experimental group (33% to 92%, McNemar $\chi^2 = 36.98$, $P < .001$). The decrease in cocaine use in the experimental treatment group was not significant (from 77% to 66%, McNemar $\chi^2 = 1.33$, $P < .3$). In the control group, neither the heroin nor cocaine use percentages changed significantly, but there was also a trend toward increased (illicit) methadone use (26% to 37%, McNemar $\chi^2 = 3.225$, $P < .10$).

To determine whether the 169 subjects included in Table 4 were a biased sample of all subjects in the study, we examined urinalysis results for 129 subjects originally assigned to the experimental treatment group and 121 subjects originally assigned to the frequent contact group for whom there was at least one follow-up urinalysis result. These represent 87% of all subjects originally assigned to the experimental treatment group and 80% of all subjects originally assigned to

TABLE 2—Selected Demographic and Drug Use Characteristics for Subjects with 1-month Follow-Up Data (N = 169)

	Experimental Treatment Group (n = 75)		Frequent Contact Group (n = 94)			
Characteristics	N	(%)	N	(%)	χ^2 or t Value	*P* Value
Sex						
Male	60	(80)	73	(78)		
Female	15	(20)	21	(22)	0.14	.71
Race						
White	11	(15)	6	(6)		
Black	19	(25)	41	(44)		
Hispanic	45	(60)	47	(50)	7.54	<.05
Currently employed	14	(19)	15	(16)	0.19	.66
Living with sexual partner	31	(41)	41	(44)	0.09	.76
Prior drug treatment	60	(80)	78	(83)	0.25	.62
Prior MMTP treatment[a]	50	(67)	51	(54)	2.67	.10
HIV status						
Positive	24	(36)	59	(67)		
Negative	43	(64)	29	(33)	14.91	<.001
Average age at baseline (yr)	33.9		35.7		t=1.71	.09
Average age began IV drug use (yr)	18.3		19.7		t=1.55	.12
Average age began IV drug use on regular basis (yr)	18.8		20.8		t=1.95	.05

[a] MMTP = methadone maintenance treatment program.

TABLE 3—Urinalysis and Self-reported Drug Use at Baseline for Subjects with 1-month Follow-up Data (N = 169)

	Experimental Treatment Group (n = 75)		Frequent Contact Group (n = 94)				
Drugs	n	(%)	n	(%)	χ^2	*P* Value	Odds Ratio (95% CI[a])
Heroin metabolites							
Urinalysis	47	(63)	58	(62)	0.02	.90	0.960 (0.513–1.795)
Heroin							
Self-report	68	(91)	89	(95)	1.02	.31	
Methadone, nonprescription							
Urinalysis	25	(33)	24	(26)	1.23	.27	0.686 (0.352–1.337)
Self-report	40	(53)	53	(56)	0.16	.70	
Heroin metabolites and/or methadone							
Urinalysis	56	(75)	66	(70)	0.41	.52	0.800 (0.404–1.583)
Heroin and/or nonprescription methadone							
Self-report	72	(96)	91	(97)	0.08	.77	
Cocaine							
Urinalysis	58	(77)	67	(71)	0.79	.37	0.727 (0.361–1.467)
Self-report	63	(84)	83	(88)	0.65	.42	

[a] CI = confidence interval.

the frequent contact control group. The urine sample collected closest to 30 days in the study was used for this analysis. For the experimental treatment group, the urinalysis at follow-up showed 47 (36%) had used heroin, 99 (77%) had used cocaine, and 117 (91%) had used methadone. For the frequent contact control subjects, 72 (60%) had used heroin, 93 (77%) had used cocaine, and 49 (40%) had used methadone. These results are similar to those reported in Table 4, although the percentage for heroin use in the experimental treatment group and the percentages for cocaine use in both the experimental and frequent contact control groups are modestly but not significantly higher. The difference in heroin metabolites between the experimental treatment group and the frequent contact control group at follow-up is still highly significant (χ^2 = 13.22, P < .001).

In order to further explore possible determinants of heroin use at 1 month, univariate analyses (chi-square tests and *t* tests) on the 169 subjects with complete intake and 1-month follow-up data were performed to determine if demographic characteristics, drug history variables, or cocaine use were associated with heroin use at 1 month. Only use of cocaine at 1 month and assignment to the control group were significantly associated with heroin use at 1 month, as presented in Table 5.

Inspection of the data showed similar relationships between heroin use and cocaine use at 1 month for both the experimental treatment group and the control group. Among the 75 subjects in the experimental treatment group, 25% had evidence of both cocaine and heroin use in their urine sample, 4% had evidence of heroin use only, 43% had evidence of cocaine use only, and 28% had no evidence of either drug (χ^2 = 4.82, P < .05). The relationship was slightly weaker among the 94 subjects in the control group: 46% had evidence of both cocaine and heroin use in their urine sample, 14% had evidence of heroin use only, 24% had evidence of cocaine use only, and 16% had no evidence of either drug (χ^2 = 2.86, P < .09). Because the urinalysis performed was capable of detecting cocaine use only for the previous several days and was capable of detecting heroin use for the previous week, these results probably underestimate concurrent use of both drugs for both groups.

Multiple logistic regression was used to examine whether treatment group status and the presence of cocaine metabolites in the 1-month urine sample were independent predictors and whether interactions between experimental group status and demographic or behavioral variables were associated with heroin use at 1 month. Such interaction effects would indicate subjects for whom interim clinic treatment might be particularly effective or not effective in reducing heroin use. The final regression equation is presented in Table 5. Both treatment group status and cocaine use at 1 month were independent predictors of heroin use at 1 month, and none of the possible interactions between treatment group and the demographic and behavioral variables were significant.

A final aspect of the experimental treatment versus control group comparisons was to examine the numbers of subjects who had entered conventional drug treatment programs at the end of data collection in June 1988, i.e., 16 months after the program began. As shown in Table 6, of the 301 subjects originally enrolled in the experimental treatment and the frequent contact control groups, 107 (72%) of those in the experimental group had been enrolled in conventional drug treatment by this data vs 85 (56%) of the controls. This difference for entering conventional treatment was statistically significant ($\chi^2 = 8.23, P < .005$).

TABLE 4—Urinalysis and Self-reported Drug Use at 1-Month Follow-up (N = 169)

	Experimental Treatment Group (n = 75)		Frequent Contact Group (n = 94)				
Drugs	n	(%)	n	(%)	χ^2	*P* Value	Odds Ratio (95% CI[a])
Heroin metabolites							
Urinalysis	22	(29)	56	(60)	15.35	<.001	3.550 (1.862–6.771)
Heroin							
Self-report	21	(28)	83	(88)	64.08	<.001	
Methadone,							
Urinalysis	69	(92)	35	(37)	52.86	<.001	0.052 (0.020–0.131)
Methadone, nonprescription							
Self-report	1	(1)	37	(39)	34.61	<.001	
Heroin metabolites and/or methadone							
Urinalysis	71	(95)	68	(72)	14.24	<.001	0.147 (0.049–0.444)
Heroin and/or nonprescription methadone							
Self-report	21	(28)	88	(94)	78.44	<.001	
Cocaine							
Urinalysis	51	(68)	66	(70)	0.09	.76	1.109 (0.575–2.138)
Self-report	29	(39)	79	(84)	37.24	<.001	

[a] CI = confidence interval.

Discussion

This study was based on the hypothesis that participation in a rapid intake, limited service methadone maintenance treatment program would reduce illicit drug use and AIDS risk behaviors among persons with a history of heroin addiction. The study attracted a group of subjects with long histories of illicit narcotic use; the great majority had previous drug abuse treatment experience. These subjects are similar in their demographic characteristics to others admitted to the Beth Israel Methadone Maintenance Treatment Program during the same time period and to other persons admitted to New York State–funded methadone treatment programs in New York City during the same time period.[8] The rate of recent cocaine use was high among these subjects, with over 60% showing cocaine metabolites in the urine specimen taken at study entry.

Random assignment studies have been difficult to conduct in the drug abuse treatment field because of difficulties in recruiting subjects into and keeping them in treatment conditions that they do not want to be in. This study experienced similar difficulties when subjects had only a one in three chance of receiving methadone treatment. After the change in the protocol, the random assignment to either immediate limited service methadone treatment or a time-limited frequent contact control group appears to have been successful. There were few difficulties in recruiting subjects with this change and there was only one variable (HIV status) on which the two groups differed. Since the actual testing of the residual serum for HIV antibody was not done until after random assignment and data collection for this study had been completed, there is no way in which the test results could have influenced the randomization, and we assume that this difference between the experimental treatment and control groups was a random effect.

Ethical considerations required that only volunteers be used in this experimental study. Thus it is not possible to ascertain how these subjects might have differed from persons on the waiting list who chose not to participate and simply waited until a regular treatment position was available. Subject recruitment varied with the perceived likelihood of receiving methadone treatment in the interim clinic, suggesting that large numbers of persons on waiting lists would apply for interim treatment if they were certain that they would receive such treatment.

Because of the substantial discrepancies in self-reported drug use at follow-up, only urinalysis results could be used as a measure of follow-up drug use. Such discrepancies are common when clients in treatment suspect that drug use will lead to some form of negative sanctions.[9] Restricting the comparisons to the urinalysis data prevented assessment of the AIDS risk associated with drug injection, but given the long histories of drug injection among these subjects, it is a reasonably safe assumption that a very high percentage of both heroin and cocaine use was by injection.

Based on the urinalysis results, participation in the interim clinic was associated with a substantial decrease in heroin use. Heroin use at 1 month was reduced by approximately half for the experimental treatment group compared to either heroin use at intake or heroin use among the control group at 1 month. Immediate intake into the limited services methadone treatment was also associated with a higher percentage of subjects being enrolled in comprehensive drug treatment. It is not surprising that receiving even limited interim services would facilitate entry into conventional treatment compared to being on a waiting list. Little is known about why many drug users apply for treatment and are placed on waiting lists but do not enter treatment, but at the least they must be considered missed opportunities for reducing illicit drug use and preventing HIV infection.

The regression analysis for possible interactions between experimental group status and other demographic and behav-

TABLE 5—Potential Predictors of Heroin in Urine Sample at the 1-Month Follow-up (n = 169)

Potential Predictors	Heroin in Urine at 30 Days N	(%)	χ^2 or t Value	*P* Value
Group				
Experimental	22	(29)		
Control	56	(60)	15.35	<.001
Sex				
Male	57	(43)		
Female	21	(58)	2.73	.10
Ethnicity				
White	5	(29)		
Black	27	(45)		
Hispanic	46	(50)	2.50	.29
Employment status at baseline				
Employed	11	(38)		
Not employed	67	(48)	1.01	.31
Living with sexual partner at baseline				
Yes	33	(46)		
No	45	(46)	0.005	.94
Previous drug treatment				
Yes	61	(44)		
No	17	(55)	1.15	.28
Previous MMTP treatment[a]				
Yes	43	(43)		
No	35	(51)	1.29	.25
HIV Status				
Positive	39	(47)		
Negative	34	(47)	0.0008	.98
Cocaine use at 30 days				
Yes	62	(53)		
No	16	(31)	7.15	<.01
Average age at baseline (yrs)	34.9		t=.03	.97
Average age began IV drug use (yrs)	19.6		t=1.07	.29
Average age began IV drug use (yrs) on regular basis	20.4		t=.91	.36

Note. In the final logistic regression equation: beta for the intercept = −0.29, with a χ^2 of 0.78, p = .3774; beta for group status = −1.30, with a χ^2 of 14.89, P = .0001; beta for cocaine in the urine = .99, with a χ^2 of 7.11, P = .008. Variables were dummy coded to 0/1. For group, a value of 1 indicates that the patient was in the treatment group. For presence/absence of drugs, a 1 indicates that the drug was present in the urine at 30 days.

[a] MMTP = methadone maintenance treatment program.

TABLE 6—Status of Subjects at End of Study (N = 301)

Status	Experimental Treatment Group (n=149) N	(%)	Frequent Contact Group (n=152) N	(%)
Lost to Contact	40	(27)	64	(42)
Drug Treatment	107	(72)	85	(56)
Jail	2	(1)	1	(1)
Death	0		2	(1)

ioral characteristics of the subjects did not produce any significant findings. Thus, the present data do not provide evidence for what types of subjects would do comparatively better or worse in interim methadone treatment. Heroin use was more common among interim treatment subjects using cocaine than those not using cocaine. Indeed, it was rare to find a treatment group subject who had evidence of heroin use but not cocaine use in the follow-up urine sample despite the greater time sensitivity for detecting heroin use. Subjects using cocaine may have used heroin simultaneously in "speedball" injections, which produce a drug effect that many users consider preferable to either drug taken alone. Cocaine users might also have taken heroin as self-medication to reduce the dysphoria following cocaine runs. Screening cocaine users out of interim methadone treatment would probably lead to greater reductions in heroin use during treatment, but would not solve the problem of what to do with the large group using both drugs.

The National Academy of Sciences[2,3] and the Presidential Commission on the HIV Epidemic[1] have recommended large-scale expansion of drug treatment in the United States as a method of reducing both illicit drug use and new HIV infections. Even if a financial commitment were made to provide large-scale expansion, it would still take a period of years for new sites to be found and for staff to be hired and trained. During this multi-year process, interim forms of treatment could be of substantial benefit as drug users waited for openings in conventional treatment. The present study did not compare interim treatment to conventional treatment, and we would object to using these findings as a rationale for substituting limited interim services for conventional treatment. We do believe, however, that these findings strongly support providing interim services rather than leaving drug users on waiting lists for conventional treatment. ☐

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The Environmental Dental Association Raises Ecological Questions as Consumer Reports Purports Mercury/Silver Fillings "Safe"

The May 1991 issue of *Consumer Reports* notes that each year up to 100 million silver-colored [amalgam] fillings are used to restore decayed teeth. The article reports that almost half of each filling is mercury. The Environmental Dental Association (EDA), a San Diego–based organization consisting of nearly 700 practitioners, maintains that the dental industry has overlooked the environmental issues associated with mercury use.

Over 100 tons of mercury are used per year in the dental industry in the United States alone. Mercury, a known poison, is more toxic than lead or arsenic. Invariably, unused amalgam material and old fillings containing mercury wind up in the hands of unregulated scrap metal dealers or in our landfills. An alarming portion of the silver/mercury filling material goes directly down the drain from dental offices, polluting our lakes, rivers, and other bodies of water. The EDA is currently initiating steps to improve on-site water filtration systems in dental offices, which would prevent further mercury contamination of water supplies.

According to the San Diego County Office of Hazardous Waste Management, the determination of hazardous waste status of dental amalgam is incumbent upon the dentist. Therefore, disposal of dental amalgam is essentially unmonitored. A spokesperson for that department stated that they have had a marked increase in calls concerning this issue, which is likely to result in stricter controls in the future.

Mike Cressler, director of Quicksilver Recycling, one of California's few treatment facilities licensed to accept mercury, recently advised the EDA that the amount of environmental contamination from dental mercury residue is significant. He expressed concern over the lack of general awareness on the part of the dental profession regarding the hazardous nature of dental waste, particularly dental amalgam. Cressler emphasized that dentists need to become more environmentally aware. He projects that the increased incidences of health and environmental liability for dentists who do not comply with environmentally-responsible practices, will likely invoke changes in the industry.

The EDA is also concerned about hazardous indoor air quality within dental offices, posing yet another risk to dental personnel who breath mercury vapor on a regular basis. The Occupational Safety and Health Administration (OSHA) estimates that 10% of all dental offices are severely mercury-contaminated. Most have adequate decontamination systems and fail to enforce the wearing of protective gear.

An FDA advisory committee met on March 15 to evaluate the use of mercury in dentistry. Although there was insufficient evidence to warrant a position that silver/mercury fillings are either safe or unsafe, two leading toxicologists testified that the issue was of legitimate concern. Research protocols to investigate health effects are currently being developed.

According to transcripts from the FDA advisory meeting, one of the world's leading authorities on mercury toxicity, Lars Friber, MD, PhD, said that "dental amalgam, from the strictly toxicological point of view, is an unsuitable dental filling material . . . steps should be taken to use, as far as possible, other material than amalgam." Contrastly, *Consumer Reports* concludes that "all risks have to be measured against the alternatives . . . amalgam fillings are still your best bet."

Although the heated debate rages on, the EDA contends that consumers of dentistry should be informed about the nature of the filling materials used and should be advised of all available options. *Consumer Reports* pointed out that the American Dental Association (ADA) opposes any informed-consent laws requiring dentists to inform patients of the possible hazards of mercury fillings and opposes discussion of alternatives.

According to the article published in *Consumer Reports*, "The question in the amalgam debate is whether the minute amount of mercury vapor thought to emanate from fillings has any health effect at all." EDA president, Joyal W. Taylor, DDS, responded to the comment as follows: "As long as this questions remains unanswered, the public deserves the right to know and the right to choose for themselves. Furthermore, we will continue to advocate environmentally responsibly dentistry."

The EDA provides scientific research data, referrals to over 700 mercury-free dentists and doctors throughout the United States and as a public service offers free information packets. The toll-free number for the EDA is 1-800-388-8124.

Public Health Briefs

Aging and the Fitness of Fire Fighters: The Complex Issues Involved in Abolishing Mandatory Retirement Ages

ABSTRACT

This study presents an assessment of the cardiorespiratory health and physical fitness of fire fighters of different ages with the aim of addressing several questions currently being debated on a national level. These include how fire fighters differ across age groups in their health and physical fitness, and what would be the impact on the existing work force of implementing minimum fitness standards as an alternative to mandatory retirement ages.

Fire fighters aged 20-65 years were found to be similar in their cardiorespiratory health and physical fitness to the sedentary segment of the general population of the same age. Particularly worrisome, however, is the low maximal aerobic capacity (31, 28, 26 ml/kg/min), high percent body fat (26, 29, 30), and high resting blood pressure (136/86, 140/90, 143/93 mm Hg) observed in the 40-45-, 50-55-, and 60-65-year-old fire fighters, respectively. In fact, 66%, 83%, and 93% of the fire fighters in these three age groups, respectively, fall below the lowest published recommendations for maximal aerobic capacity in this profession.

These data underscore the need to establish minimum health and physical performance standards for fire fighters and demonstrate the profound impact such minimum standards will have on the existing work force. (*Am J Public Health*. 1991;81:1192–1194)

Kurt Saupe, MS, Mark Sothmann, PhD, and Deborah Jasenof, MS

Introduction

The feasibility of abolishing mandatory retirement ages in public safety occupations is being studied under US congressional mandate. One suggestion has been to replace mandatory retirement ages with minimum standards for health and physical fitness.[1] Before serious discussion regarding developing and implementing standards can begin, the impact of those standards on the current work force must be known. Therefore, we conducted a descriptive case study of a large municipal fire department (approximately 5000 uniformed personnel) in order to (1) describe differences across age groups in selected measures of physical fitness and health using a cross-sectional design and (2) develop hypotheses on the impact of implementing physical performance standards on the existing work force

Methods

Subjects were recruited from a randomized listing of all uniformed full-time fire fighters from a large municipal fire department. The 150 subjects were divided by age into five groups of 30 each: aged 20 to 25, 30 to 35, 40 to 45, 50 to 55, and 60 to 65 years. No more than 40 individuals in each age group had to be contacted in order to obtain the 150 volunteers.

Cardiovascular status was determined from a graded exercise test employing a modified Balke treadmill protocol. ECG, blood pressure, and heart rate were taken at rest; at the end of each stage of exercise; and at 1, 3, and 5 minutes during recovery. Expired gases were collected during the last 1 to 4 minutes prior to volitional fatigue or symptom-limited test termination. Maximal oxygen uptake (VO2max) was measured using an open circuit system as described elsewhere.[2]

Pulmonary function testing was performed using a Collins Survey Spirometer. From a forced vital capacity (FVC) maneuver, forced expiratory volume at 1 second (FEV1) and forced expiratory flow rate between 25% and 75% of FVC (FEF25–75) were calculated using standard procedures.

Percentage of body fat was estimated based upon skinfold thickness of the abdomen, chest, and thigh as measured with a Lange skinfold caliper. The average of three measurements at each site was used in a regression equation to predict percentage of body fat.[3]

Measurements of grip strength, standing long-jump, and situps were made for inclusion with selected cardiovascular and anthropometric variables in the composite fitness evaluation system developed by Davis et al.[4]

Analysis of variance using a post hoc Scheffe test was employed to determine any statistically significant ($P < .05$) differences between the age groups.

Results

Aerobic capacity significantly decreased from the 20- to 25-year-old group (47.7 ml/kg/min) to the 30- to 35-year-old group (37.9 ml/kg/min) and from the 30- to

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At the time of this study, all three authors were with the Department of Human Kinetics, University of Wisconsin, Milwaukee. Kurt Saupe is now with the Physiology Department of the University of Wisconsin, Madison.

Requests for reprints should be sent to Kurt Saupe, 504 N Walnut Street, University of Wisconsin, Madison, WI 53705.

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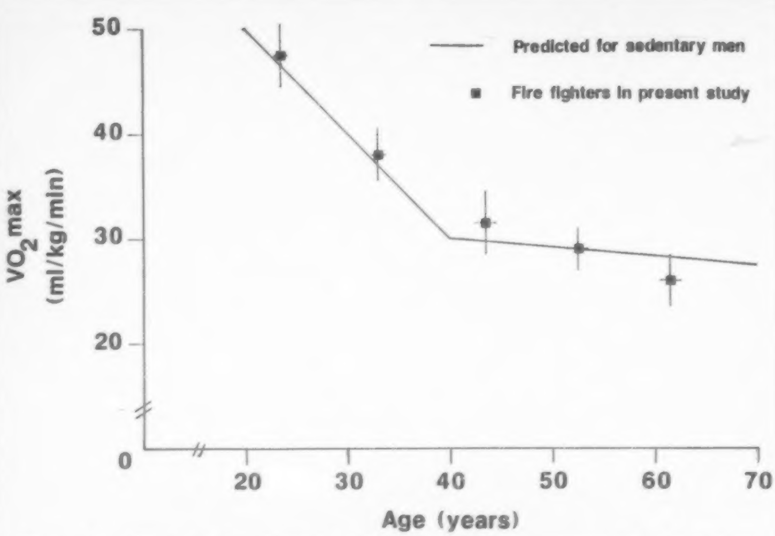

FIGURE 1—Decline in VO2max with age in fire fighters compared with the decline predicted for sedentary individuals by Buskirk and Hodgson.[17]

TABLE 1—Percentage of Fire Fighters above Proposed Minimum Values for VO2max

	Age Groups[a]					
Proposed Values	20–25	30–35	40–45	50–55	60–65	Reference
3.0 l/min	83	73	27	33	3	13,14
42 ml/kg/min	87	20	3	0	0	15
39 ml/kg/min	93	40	13	3	3	16
33.5 ml/kg/min	100	83	33	17	7	5

[a] n = 30 for each group.

TABLE 2—Characteristics of Fire Fighters across Age Groups[a]

	Age Groups[a]				
Characteristics	20–25	30–35	40–45	50–55	60–65
Resting systolic blood pressure (mm Hg)[b]	128 ± 10	133 ± 7	137 ± 14	141 ± 15	142 ± 15
Resting diastolic blood pressure (mm Hg)[b]	72 ± 12	82 ± 9	86 ± 9	88 ± 11	88 ± 7
Forced vital capacity (l BTPS)[b]	4.99 ± 0.5	5.30 ± 0.6	4.76 ± 0.7	4.53 ± 0.6	3.80 ± 0.6
FEVI (l BTPS)[b]	4.19 ± 0.4	4.26 ± 0.5	3.70 ± 0.7	3.62 ± 0.5	2.82 ± 0.7
FEF25–75 (l/s)[b]	4.69 ± 0.8	4.43 ± 1.0	3.47 ± 1.1	3.80 ± 1.1	2.32 ± 1.1
% Smokers	10	33	33	27	10
% Previous smokers	10	27	43	50	80
% Never smoked	80	40	23	23	10
% Who exercise more than twice per week	60	40	17	10	33

Note: BTPS = body temperature and pressure, saturated; FEV1 = forced expiratory volume at 1 second; FEF25–75 (l/s) = forced expiratory flow rate between 25% and 75% of forced vital capacity.
[a] n = 30 for each group.
[b] Data are expressed as mean ± standard deviation.

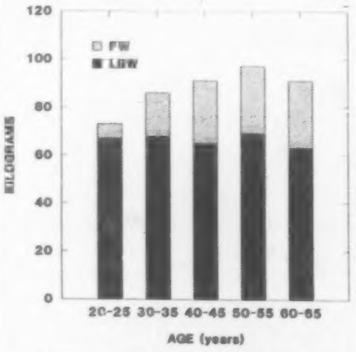

FIGURE 2—Lean body weight (LBW) and fat weight (FW) as a portion of total body weight in fire fighters across age groups.

35-year-old group to the 40- to 45-year-old group (31.5 ml/kg/min) (Figure 1). Given these values for VO2max, Table 1 indicates the percentage of fire fighters in each of the five age groups who would pass various published recommendations for VO2max in fire fighters.

Resting data on blood pressure and pulmonary function are presented for each of the age groups in Table 2. Both systolic and diastolic pressures are significantly higher in the 50- to 55-year-olds and 60- to 65-year-olds than in the 20- to 25-year-olds. FVC was significantly less in the 60- to 65-year-olds than in the 20- to 25-year-old group. Also presented in Table 2 are self-reported data on smoking history and physical activity.

Figure 2 presents lean body weight and fat weight across age groups. Percentage of fat increased significantly from ages 20 to 25 years (8.3%) to 30 to 35 years (19.4%) and again from 30 to 35 years to 40 to 45 years (27.2%).

Table 3 indicates the percentage of fire fighters in each age group exceeding certain clinical criteria. Fire fighters with over 25% body fat are considered to be moderately obese.[5] Individuals with a resting diastolic pressure (DBP) between 90 and 114 mm Hg or a systolic pressure (SBP) greater than 140 mm Hg are considered to the borderline to moderately hypertensive[6] and are categorized as such in Table 3. The percentage of fire fighters in each age group manifesting positive stress tests according to criteria established by the American College of Sports Medicine[7] are listed in Table 3. Exercise hypertension (SBP > 250 or DBP > 120) was the most frequently observed problem in fire fighters aged 40 to 55. In con-

TABLE 3—Percentage of Fire Fighters in Each Age Group Exceeding Selected Clinical Criteria

	Age Groups[a]				
Criteria	20–25	30–35	40–45	50–55	60–65
> 25% Body Fat	0	27	60	83	77
Borderline to moderate hypertension	20	33	53	70	73
Positive exercise stress test	0	0	7	10	40
Mild to moderate pulmonary impairment	10	13	30	23	53

[a] n = 30 for each group

trast, 10 of the 12 positive stress test results in the 60- to 65-year-old age group were due to ECG abnormalities. The average VO2max of the fire fighters with positive stress tests was 24.5 ml/kg/min. Individuals with either an FEV1/FVC of less than .74 or an FVC less than 79% of predicted[8] are considered to have mild to moderate pulmonary impairment[9] and are also presented in Table 3.

A test battery for assessing the overall physical fitness of fire fighters, which includes measures of strength, cardiovascular status, and pulmonary function and has been shown to be predictive of performance time on a standardized fire suppression test,[4] was also used. According to this rating system, 38% of fire fighters aged 40 to 45, 60% aged 50 to 55, and 96% aged 60 to 65 would be considered poor in their overall fitness for fire fighting.

When fire fighters in each age group were divided into those who reported exercising more than twice a week and those exercising less than twice a week, no significant differences in VO2max, percentage of body fat, or FVC were observed.

Discussion

This case study of a large metropolitan fire department illustrates the complexity of the mandatory retirement age issue faced by municipalities. These data suggest that substantial age-associated declines do occur in the health and physical fitness of fire fighters, although these declines are not necessarily age-dependent. Although fitness standards have been advocated as an alternative to mandatory retirement ages, the data presented here demonstrate that using published recommendations as standards[2,10,13] for the aerobic fitness of fire fighters would have an "adverse impact"[14] on a large segment of the existing work force. Given the existing legal controversies pertaining to adverse impact, a municipality would have to present considerable justification for any proposed standard, which would involve costly research and litigation.[15,16] Lowering standards to avoid adverse impact is legitimate only if public safety is not compromised. Our data suggest that any consideration of a shift in national policy away from mandatory retirement ages should be accompanied by an in-depth examination of the feasibility of developing and implementing medical and fitness standards at the municipal level, and clarification of the controversial issue of whether fitness standards that have adverse impact on certain protected groups are necessary to maintain public safety. ☐

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Malnutrition in Elderly Ambulatory Medical Patients

Aaron Manson, MD, and Steven Shea, MD

ABSTRACT

Elderly ambulatory persons may be especially susceptible to malnutrition, particularly those who are poor and socially isolated or have comorbid chronic medical diseases. We found that 98 of 2,986 persons aged 60 years or older attending a hospital-based medical practice between 1979 and 1989 weighed less than 45.4 kg (100 lbs). All but 1 of these subjects met criteria for malnutrition as judged against age-specific norms for weight. Thus the prevalence of malnutrition in this sample was 3.25% (95% CI 2.61, 3.89%). Interviews and physical examinations of a subsample (n = 16) revealed that all 16 subjects either met anthropometric-based criteria for malnutrition or were being treated for malnutrition. Of the 98 subjects who weighed less than 45.4 kg, 62 (63.3%; 95% CI 53.8, 72.8%) had comorbid conditions that could have contributed to malnutrition. Physicians did not record a diagnosis of malnutrition or weight loss in 47.9% of subjects (95% CI 38.0, 57.8%) and did not prescribe a nutrition supplement for 76.5% of subjects (95% CI 68.1, 84.9%). Subjects treated with nutrition supplement were more likely to have cancer. These findings suggest that malnutrition, both with and without concomitant major comorbid disease, is relatively frequent among elderly ambulatory patients and that a specific nutritional diagnosis is not made in many cases. We suggest that weight under 45.4 kg in an elderly person is a useful criterion for identifying elderly patients at nutritional risk. (*Am J Public Health.* 1991;81:1195–1197)

Introduction

In 1968 Jean Mayer testified before Congress that as many as 10% of Americans were malnourished due to poverty.[1] During the ensuing decade, government programs reduced malnutrition, but more recently funding for these programs has diminished.[2] The two largest nutritional surveys in the United States, the Ten-State Nutrition Survey[3] and the National Health and Nutrition Examination Surveys (NHANES I and II),[4–6] collected comprehensive data but were conducted in the 1970s, and neither study estimated the prevalence of malnutrition in the elderly. Other studies of malnutrition or weight loss among the elderly are anecdotal, describe small numbers, include hospitalized patients, or do not estimate the prevalence of malnutrition in the populations studied.[7–24] We therefore studied the prevalence of malnutrition among elderly ambulatory patients.

Methods

The study population consisted of outpatients aged 60 years or older, who were receiving care in the medical faculty practice at Columbia-Presbyterian Medical Center, and whose recorded weights in the practice database were 45.4 kg (100 lbs) or less on at least two office visits during the period from 1979 to 1989. Patients were routinely weighed on a balance scale at most visits. In 1989, there were 8,956 patient records in the database; of these 2,987 (33.3%) referred to patients who were age-eligible; one age-eligible patient did not have a recorded weight in the database. Hospital charts of 124 of the 128 age- and weight-eligible patients were obtained, and these patients' weight of 45.4 kg or less on at least two visits in 98 of these 124 patients was confirmed (Table 1).

We reviewed the medical records of these 98 patients using standardized data forms. The diagnoses of malnutrition or weight loss were considered to have been recorded if there was explicit mention in the chart subsequent to the index visit. Treatment of weight loss or malnutrition was considered to have been provided if chart notes recorded the prescription of liquid nutrition supplement. Alcoholism, cancer, severe chronic pulmonary disease, congestive heart failure, dementia, depression, or stroke were noted as comorbid conditions that could contribute to malnutrition.

We further defined nutritional status on the basis of two sets of age- and sex-specific norms, NHANES[25] and the Ten-State Nutrition Survey.[3] Subjects were considered severely malnourished if measurements of weight for height[25] or weight[3] were below the fifth percentile according to these norms. Subjects were considered moderately malnourished if these measurements were between the fifth and fifteenth percentiles.

Results

Laboratory findings in the 98 subjects were consistent with although not specific for malnutrition (Table 1). Of the 48 subjects with recorded heights and weights, 28 (58.3%) were severely malnourished, and 15 (31.3%) were moderately malnourished as measured against the norms from NHANES. All five of the remaining subjects, who could not be classified because their heights were below the minimum height in these distributions, either were taking nutrition supplements or had a diagnosis of malnutrition in the medical record. Measured against norms based on weight alone[3] 2 subjects more than 85 years old could not be classified, 48 (49.0%) were severely malnourished, 42 (42.9%) were moderately malnourished, and 6 (6.1%) were above the 15th percentile. Of the 6 subjects above the 15th percentile, 4 were described as malnourished

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Aaron Manson is with the Division of General Medicine, Department of Medicine. Steven Shea is with the Division of General Medicine, Department of Medicine, and the Division of Epidemiology, School of Public Health, Columbia University.

Requests for reprints should be sent to Dr. Manson at Columbia-Presbyterian Medical Center, VC 2-205, 622 West 168th Street, New York, NY 10032.

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TABLE 1—Characteristics of 98 Subjects with Weights of 45.4 kg (100 lbs) or Less Attending a General Medicine Ambulatory Care Practice in New York City, 1979–1989[a]

	Mean		SD
Age (years)	75.2	±	9.8
Follow-up time (months)	33.1	±	32.4
Weight (kg)	40.9	±	3.9 (range 26.1–45.4)
Height (cm)	155.4	±	10.1 (range 134.8–175.3)
Albumin (mg/dl)	4.0	±	0.46
Hemoglobin (mg/dl)	12.3	±	1.8
Absolute lymphocyte count (cells/microliter)	1,647	±	761

[a] Five (5.1%) of the subjects were men, and 28 (28.6%) were receiving Medicaid.

TABLE 2—Characteristics of 98 Subjects with Weights of 45.4 kg (100 lbs) or Less Classified by Nutritional Diagnosis as Recorded in the Medical Record

Basic Characteristics	Subjects with Diagnosis of Malnutrition or Weight Loss (n = 51) Mean SD	Subjects with No Nutritional Diagnosis (n = 47) Mean SD
Age (years)	75.6 ± 9.4	74.7 ± 10.4
Follow-up time (months)	29.9 ± 27.4	36.5 ± 37.0
Weight (kg)[a]	40.8 ± 4.0	41.1 ± 3.9
Height (cm)	153.3 ± 11.2	157.8 ± 8.2
Albumin (mg/dl)[b]	3.97 ± 0.42	4.04 ± 0.51
Hemoglobin (mg/dl)[c]	12.2 ± 1.6	12.4 ± 2.0
Absolute lymphocyte count[d] (cells/microliter)	1,767 ± 816	1,519 ± 690
Number (percent) men	4 (7.8%)	1 (2.1%)
Comorbid Disease[e]	**Subjects (Percent) with Diagnosis Malnutrition or Weight Loss (n = 51)**	**Subjects (Percent) with No Nutritional Diagnosis (n = 47)**
Alcoholism	5 (9.8%)	3 (6.4%)
Cancer	11 (21.6%)	4 (8.5%)
Chronic pulmonary disease	8 (15.7%)	15 (31.9%)
Depression	10 (19.6%)	9 (19.1%)
Congestive heart failure	9 (17.6%)	11 (23.4%)
Dementia	9 (17.6%)	5 (10.6%)
Stroke	3 (5.9%)	5 (10.6%)

[a] Height was available for 48 subjects.
[b] Serum albumin was available for 80 subjects.
[c] Hemoglobin was available for 77 subjects.
[d] Absolute lymphocyte count was available for 60 subjects.
[e] Sixty-two subjects had one or more comorbid conditions; 28 subjects had more than one.

in the medical record, and a fifth subject met criteria for malnutrition as measured against norms of weight for height.[25] Both extremely elderly subjects were described as malnourished in the medical record. Thus the false positive rate in classifying these 98 subjects using weight alone was 1% or less. We therefore estimated the prevalence of malnutrition among subjects over the age of 60 years in this medical practice as 97 of 2,986 (3.25%; 95% CI 2.61, 3.89%).

The diagnosis of malnutrition was recorded in the charts of 23 subjects (23.5%), and the diagnosis of weight loss was recorded in 28 additional subjects (28.6%, Table 2). Twenty-three subjects (23.5%) were treated with calorie supplement (Table 3). Thus 47 of 98 subjects (47.9%; 95% CI 38.0, 57.8%) had no recorded diagnosis of malnutrition or weight loss, and 75 (76.5%; 95% CI 68.1, 84.9%) were not treated with calorie supplement.

Comorbid conditions that could have contributed to malnutrition were present in 62 subjects (63.3%; 95% CI 53.8, 72.8%, Table 2). Presence of a comorbid condition was not significantly associated with a recorded diagnosis of malnutrition or weight loss; however, such subjects were more likely to have cancer (odds ratio = 2.96; 95% CI 0.87, 10.04).

A substudy (N = 16) of subjects who made scheduled visits to the practice between July 1989 and February 1990 found that the observed weight was 45.4 kg or less in all subjects and that the mean observed weight differed from that in the chart by less than 1%. Using norms for bone-free upper arm muscle area based on triceps skinfold and mid-arm circumference measurements,[25] we found 11 subjects to be below the fifteenth percentile. Of the remaining 5, 4 had a recorded diagnosis of malnutrition or were taking nutrition supplement. The remaining subject met criteria for malnutrition according to norms for weight[3] and weight for height.[25] Thus, all 16 substudy subjects either met anthropometric criteria for malnutrition or were being treated for malnutrition, including 7 subjects with no recorded nutritional diagnosis.

Discussion

This study found that 3.25% of ambulatory elderly patients in a medical practice had evidence of malnutrition, that weight of 45.4 kg (100 lbs) or less is a useful screening criterion for malnutrition, and that the diagnosis of malnutrition was noted in only 52.1% of the 98 malnourished subjects. Of the 98 subjects, 62 (63.3%) had an underlying disease that probably contributed to malnutrition, but more than a third did not; this suggests that malnutrition among the elderly is often present in the absence of major comorbid illness.

It is likely that our method underestimated the prevalence of malnutrition. A weight of 45.4 kg may be too low to identify some malnourished men with larger frames. Thus, only 5 subjects in this series were male.

Although data were collected retrospectively, it is unlikely that information regarding other diagnoses was incomplete since there were computer-based lists of medical diagnoses for all subjects, and the mean observation time in the medical record was 33 months. Weights recorded in the chart were measured during routine care; however, weights measured in the validation substudy corroborated those in the chart.

Although their findings are not conclusive, several longitudinal studies have found that extreme leanness is associated with adverse medical outcomes.[26,27] Al-

TABLE 3—Characteristics of 98 Subjects[a] with Weights of 45.4 kg (100 lbs) or Less Classified by Nutritional Treatment as Recorded in the Medical Record[b]

	Treated (n = 23)		Not Treated (n = 75)	
Men	Mean	SD	Mean	SD
Age (years)	75.0	± 8.7	75.2	± 10.2
Follow-up time (months)*	18.7	± 19.7	37.4	± 34.5
Weight (kg)	39.9	± 4.1	41.2	± 3.8
Height (cm)	154.3	± 11.1	155.9	± 9.6
Albumin (mg/dl)	3.80	± 0.41	4.07	± 0.46
Hemoglobin (mg/dl)**	12.5	± 1.0	12.2	± 2.0
Absolute lymphocyte count (cells/microliter)	1,403	± 567	1,720	± 802

[a]Four (17.4%) of the treated subjects were men, and 7 (30.4%) had cancer.
[b]Fisher's Exact Test was used to determine statistical significance.
*P = .002. **P = .003.

though the prognostic significance of a weight of 45.4 kg or less (100 lbs) in the elderly is not known, our data indicate that this easily remembered indicator is a useful screening criterion for malnutrition in the elderly and can be used to define the prevalence of this condition.[28] ☐

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Suicides among Older United States Residents: Epidemiologic Characteristics and Trends

Patrick J. Meehan, MD, Linda E. Saltzman, PhD, and Richard W. Sattin, MD

ABSTRACT

Suicide rates for elderly US residents decreased between 1950 and 1980, but have increased recently. We analyzed suicide mortality trends using national mortality data for the period 1980 through 1986. Suicide rates during this period increased for each 5-year age group over age 65. Elderly White males have the highest suicide rates and experienced a rate increase of 23%. The rate for Black males rose by 42%. Divorced males have the highest age-adjusted sex- and marital status-specific rates, and experienced a rate increase of 38% over the 7-year period. Suicide rates among older US residents vary by region of the country and are highest in the West. Rates increased in all regions except the Northeast. Firearms are the most common method of suicide in the elderly, and firearm use increased during this period from 60% to 66% of all suicides. Given the recent increase in suicide rates for the elderly and the magnitude of the problem in this age group, it is again important to direct our attention to the problem of suicide in the elderly and recognize the need for effective prevention strategies. (*Am J Public Health.* 1991;81:1198–1200)

Introduction

Suicide rates for elderly US residents decreased between 1950 and 1980,[1] but have increased recently.[2] Previous studies of suicide among persons of all ages have identified high risk groups including whites, males, widowed or divorced persons, and residents of the western United States.[1,3,4] In this article we examine recent trends in suicide rates for elderly US residents, with an analysis of risk groups and methods used in suicide.

Methods

Data on suicide deaths for the years 1980 through 1986 are from annual mortality data tapes compiled by the National Center for Health Statistics (NCHS), Centers for Disease Control. Suicides analyzed in this article are deaths among persons over age 65 with external cause-of-death codes E950–E959, according to the International Classification of Diseases, Ninth Revision (ICD-9).

Population data used in the calculation of rates by age, sex, and race are from annual computer data tapes produced by the US Bureau of the Census. Marital status-specific and region-specific rates were calculated using population estimates from Bureau of the Census publications.[5,6] Marital status-specific rates were age-adjusted by the direct method (10-year age groups), using the 1980 population as the reference. Because of the small numbers of suicides among other racial groups, we limited our analysis of race-specific rates to the Black and White races.

Results

From 1980 through 1986, 36 789 suicides were reported among US residents over age 65 years, and crude annual suicide rates increased for this age group by 21% (from 17.8 to 21.5 per 100 000). Rate increases ranged from 9% (from 16.2 to 17.7 per 100 000) for the 65- through 69-year age group to 38% (from 18.0 to 24.8 per 100 000) for the 80- through 84-year age group (Table 1). During this period, men accounted for 80% of the suicides among persons over age 65, and rates for men increased for each 5-year age group over age 65.

White males had the highest race- and sex-specific suicide rates for persons over age 65 years (Table 2), and from 1980 to 1986 their rate increased 23% (from 37.2 to 45.6 per 100 000). Suicide rates for Black men increased 42% (from 11.4 to 16.2 per 100 000), and the suicide rates for White women increased 17% (from 6.4 to 7.5 per 100 000).

The 1986 suicide rate for divorced males over age 65 was 3.2 times the rate for married men and 18.9 times the rate for married women (Table 3). From 1980 to 1986, the rate for divorced men increased 38% (from 79.5 to 109.7 per 100 000). Rates for women increased, regardless of the marital status category. However, because the numbers are small, the rates fluctuate more.

In 1986, the western United States had the highest suicide rate for persons over age 65 (29.5 per 100 000), the Northeast had the lowest rate (13.6 per 100 000), and the southern and northcentral regions had intermediate rates (23.7 and 20.2 per 100 000, respectively). From 1980 to 1986, the suicide rate decreased slightly in the Northeast, whereas rates increased in the northcentral (22%), western (23%), and southern (29%) regions.

From 1980 to 1986 firearms were the most common method of suicide for both men and women over age 65 years (Figure 1). Over the 7 years, 73% of older men and 29% of older women committed suicide with firearms. Firearm use increased from

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The authors are with the National Center for Environmental Health and Injury Control, Centers for Disease Control, Atlanta, Ga.

Requests for reprints should be sent to Mailstop F36, Intentional Injuries Section, Centers for Disease Control, 1600 Clifton Road, NE, Atlanta, GA 30333.

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60% of all over-65 suicides in 1980 to 66% in 1986 for both sexes combined.

Discussion

Suicide is the third leading cause of deaths from injury among older US residents, following deaths from falls and from motor vehicle crashes.[2] From 1980 through 1986, suicide rates increased for each 5-year age group over age 65; the only other age groups that experienced a rate increase of at least 5% over this period were the 60- through 64-year age group (10%) and the 15- through 19-year age group (20%) (NCHS Detailed Mortality Tapes 1980–1986).

Although the demographic groups at high risk for suicide among the elderly are similar to those for all ages combined,[1,3,4] differences in suicide rates between demographic subgroups have become more pronounced. For example, the rate ratio between men and women in the >85 age group increased from 9.2 to 13, and the rate ratio for the western region compared with the northeast region increased from 1.6 to 2.2. It is noteworthy, however, that from 1980 to 1986 the rate ratio for Whites over age 65 compared with Blacks narrowed, due to the larger increase in suicide rates among Black men.

In 1986, 66% of suicides among the elderly were committed with a firearm, compared with 57% among persons under age 65. Methods of suicide differ more between men and women over age 65 than for other age groups. In 1986, 75% of male victims over age 65 died of gunshot wounds, compared with 61% of male victims under age 65. During the same year, 31% of female victims over age 65 died of gunshot wounds, compared with 41% under age 65 (NCHS Detailed Mortality Tapes 1986). The use of drugs and poisons is a common method of suicide for older women. Not only do methods of suicide have important implications for the care of acutely suicidal persons, but the availability of specific methods as risk factors for suicide merits further study.

Possible limitations of our data should be discussed. Underreporting has been suggested to be a major source of error in the determination of suicide rates.[4,7] Although the precise extent to which deaths are underreported as suicides is unknown, it is probably a small proportion and should not affect the analysis of trends in suicide rates.[8,9] The denominator used in the calculation of marital status-specific rates is derived from the current population survey, and excludes institutionalized persons. Other rates were calculated using census data and intercensal estimates of the resident US population. Although this methodological difference may affect actual rates, any effect on trends in age-adjusted rates should be minimal.

TABLE 1—Numbers of Suicides and Suicide Rates by Sex and Age Group for US Residents over Age 65, 1980–1986ᵃ

		Age Group									
		65–69		70–74		75–79		80–84		>85	
Year	Sex	No.	Rate	No.	Rate	No.	Rate	No.	Rate	No.	Rate
1980	Male	1098	28.1	955	33.2	766	41.3	446	43.3	345	49.9
	Female	324	6.6	253	6.4	175	5.9	90	4.6	85	5.4
	Total	1422	16.2	1208	17.7	941	19.5	536	18.0	430	18.9
1981	Male	1033	26.0	925	31.4	742	38.8	484	46.2	354	50.4
	Female	356	7.2	260	6.4	168	5.5	93	4.7	63	3.8
	Total	1389	15.6	1185	16.9	910	18.4	577	19.1	417	17.7
1982	Male	1129	28.0	1053	34.9	859	43.7	518	48.3	363	50.4
	Female	361	7.2	269	6.5	197	6.3	104	5.2	66	3.8
	Total	1490	16.5	1322	18.5	1056	20.8	622	20.1	429	17.5
1983	Male	1139	27.7	1108	35.9	913	45.2	622	56.7	397	53.7
	Female	370	7.3	307	7.3	210	6.6	126	6.1	92	5.1
	Total	1509	16.4	1415	19.4	1123	21.5	748	23.6	489	19.3
1984	Male	1212	29.1	1239	39.3	939	45.2	614	54.7	390	51.6
	Female	387	7.6	303	7.1	208	6.3	133	6.3	92	4.9
	Total	1599	17.2	1542	20.7	1147	21.4	747	23.0	482	18.4
1985	Male	1236	29.1	1252	39.0	1035	48.5	714	61.9	428	55.4
	Female	345	6.7	312	7.2	247	7.4	130	6.0	89	4.6
	Total	1581	16.8	1564	20.7	1282	23.3	844	25.4	517	19.1
1986	Male	1344	30.7	1368	41.9	1161	52.8	695	58.5	484	61.1
	Female	363	6.9	339	7.7	272	7.9	155	6.9	94	4.7
	Total	1707	17.7	1707	22.3	1433	25.5	850	24.8	578	20.7

ᵃPer 100 000 population per year.

TABLE 2—Numbers of Suicides and Suicide Rates by Race for US Residents over Age 65, 1980–1986ᵃ

	White				Black			
	Male		Female		Male		Female	
Year	No.	Rate	No.	Rate	No.	Rate	No.	Rate
1980	3490	37.2	896	6.4	97	11.4	18	1.4
1981	3407	35.7	899	6.3	105	12.1	30	2.3
1982	3786	38.8	961	6.6	108	12.2	23	1.8
1983	4020	40.3	1068	7.2	128	14.1	19	1.4
1984	4233	41.6	1081	7.2	129	14.0	24	1.7
1985	4495	43.2	1059	6.9	143	15.2	38	2.7
1986	4853	45.6	1168	7.5	156	16.2	35	2.4

ᵃPer 100 000 population per year.

Known risk factors for suicide that may contribute to high rates and to an increase in rates for the elderly include depression, alcoholism, and chronic illness.[10] Although overall rates of depression may be increasing in the United States,[11] this increase would have to be occurring disproportionately among the elderly to explain the prominent and largely isolated increase in suicide rates for this group. We are not aware of such evidence. Alcohol abuse is less prevalent among the elderly than in younger age groups.[12] It is possible, however, that more people with alcohol-related illnesses are living to an older age, resulting in an increase in the prevalence of alcoholism among persons over age 65 and therefore more at-risk people in this age group. Fi-

TABLE 3—Age-Adjusted Suicide Rates by Sex and Marital Status for US Residents over Age 65, 1980–1986*

	Marital Status							
	Married		Never Married		Divorced		Widowed	
Year	Male	Female	Male	Female	Male	Female	Male	Female
1980	26.9	5.0	61.6	4.9	79.5	10.8	76.5	7.7
1981	27.0	5.3	55.3	5.8	74.7	10.3	69.5	7.3
1982	27.5	5.4	74.6	5.9	99.9	16.7	81.5	7.1
1983	29.9	5.7	69.8	6.8	103.4	13.2	76.5	8.0
1984	31.0	5.6	65.9	7.5	113.6	18.3	81.4	8.2
1985	33.1	5.5	53.5	6.1	101.1	15.0	78.7	8.0
1986	34.5	5.8	53.5	5.9	109.7	18.4	85.9	8.4

*Per 100 000 population per year.

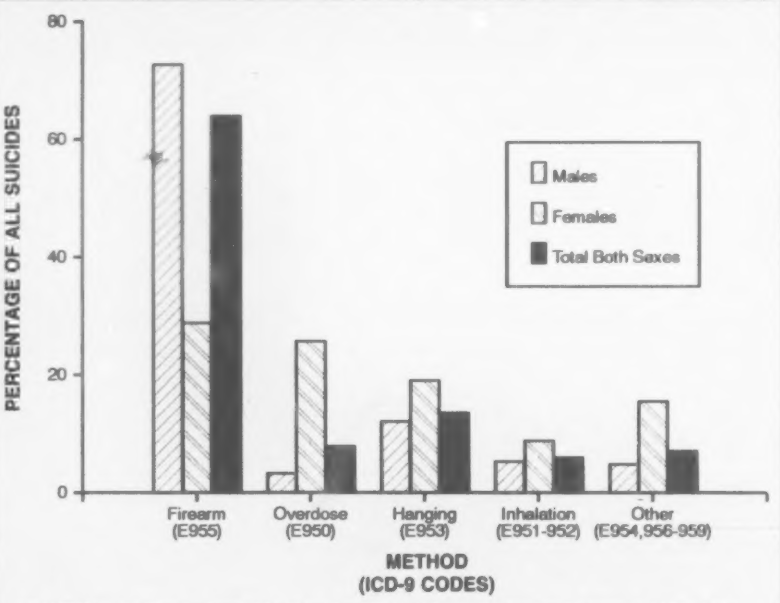

FIGURE 1—Method of suicide used by US residents over age 65 years, 1980–1986

nally, advances in medical care are helping to prolong the lives of persons with chronic illnesses. This could result in higher suicide rates for the elderly, because suicide risk may be higher among chronically ill persons.[13]

Several recommendations can be made based on our findings. We need to consider how current approaches to suicide prevention can better reflect the special circumstances of older persons. Crisis services, for example, may have the greatest impact on the prevention of youth suicide, which may be more impulsive in nature.[14] For older adults, however, addressing the increasingly prevalent problem of social isolation,[15] perhaps reflected in the high suicide rates among divorced persons, may be a more effective prevention strategy. Studies that address the role of economic factors in the precipitation of suicide are needed, as are analytic studies to evaluate, on the individual level, the relationship between correlates of social disruption, such as divorce or geographic migration, and the risk of suicide. Finally, health providers should be increasingly aware of the potential for suicide among the elderly, particularly among those at highest risk, such as divorced older men and Black men. ☐

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Development and Validation of a Simple Device to Estimate Birthweight and Screen for Low Birthweight in Developing Countries

Nikorn Dusitsin, MD, MS, Sumana Chompootaweep, MD, MS, Piyalamporn Poomsuwan, MS, Kanchana Dusitsin, MD, Prasertsri Sentrakul, MD, MS, and Pisake Lumbiganond, MD, MS

ABSTRACT

Low birthweight (<2500 g) is the major factor associated with the death of infants within the first 4 weeks of life. The mean birthweight of newborn babies of a certain geographic area reflects the quality of maternal and child health care service as well as the degree of socioeconomic development of that particular region. Hence, birthweight is being used increasingly as an indicator for health and socioeconomic planning. However, in a developing country such as Thailand, two-fifths of the babies are delivered at home and are not weighed because scales are not available. To solve this problem in rural areas, a circular nomographic chart was developed with which the birthweight can be computed from a newborn baby's chest and mid-arm circumferences. Preliminary trials comparing these charts with standard baby scales showed a high degree of accuracy with sufficient sensitivity and specificity. (*Am J Public Health*. 1991;81:1201–1205)

Introduction

Low-birthweight (<2500 g) is the major factor associated with the death of infants within the first 4 weeks of life. The mean birthweight of newborn babies of a certain geographic area may reflect the quality of maternal and child health care service as well as the degree of socioeconomic development of that particular region.[1–4] However, in a developing country such as Thailand, two fifths of the babies are delivered outside hospitals and are not weighed because scales are not available.[5] Anthropometric parameters such as chest circumference,[6] mid-arm measurement,[7] and foot length[8] have been used to identify low birthweight and a newborn at risk. The World Health Organization has studied the chest and arm circumferences in correlation with birthweight, aiming to develop chest or arm bands that would indicate that the birthweights are above or below a specified cutoff point (i.e., 2500 g). However, no effort has been made to correlate anthropometric parameters with actual birthweight. This study uses multiple parameters of chest and arm circumferences and the length of a newborn to estimate birthweight and to validate this new method in various centers throughout Thailand.

Methods

The Training Data Set

The training data set undertaken at the Department of Pediatrics, Lerdsin Hospital, from January to March 1985, involved 402 newborn babies, 197 male and 205 female. The duration of pregnancy at the time of delivery ranged from 36 to 40 weeks according to Nägele's formula. The measurements taken within 24 hours of birth included chest circumference at nipple line in centimeters, arm circumference, midway between the acromion process and the tip of elbow in centimeters, crown-to-heel length in centimeters, and weight in grams.

Measuring Instruments

The measuring instruments consisted of a tailor's metric measuring tape, tested against a standard ruler, and a baby metric beam balance (SECA, made in Germany), tested for accuracy with standard weights. All measurements were performed by the same pediatrician throughout the study.

Sample Size

The sample size was calculated after Lachin[9] to estimate the sample size for the correlation (r) .85, setting $\alpha = .05$ and the power of test at .80 (for a two-tailed test).

Data Analysis

Stepwise multiple regression analysis was employed to formulate the best linear prediction equation.[10] The least squares method was used for fitting the "best" straight line to a given set of data in order to relate two variables.

Testing Data Set

The testing data set of body weight and chest and arm circumferences of 3222 newborn infants from eight hospitals in north, northeast, central, and south Thai-

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Nikorn Dusitsin and Prasertsri Sentrakul are with the Faculty of Medicine, Chulalongkorn University, Bangkok, Thailand. Sumana Chompootaweep and Piyalamporn Poomsuwan are with the Institute of Health Research, Chulalongkorn University. Kanchana Dusitsin is with Lerdsin Hospital, Ministry of Public Health, Bangkok. Pisake Lumbiganond is with the Faculty of Medicine, Khon-Kaen University, Thailand.

Requests for reprints should be sent to Nikorn Dusitsin, MD, Department of Obstetrics & Gynaecology, Faculty of Medicine, Chulalongkorn University, Rama IV Road, Bangkok 10330, Thailand.

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TABLE 1—Mean ± Standard Deviation for Weight, Chest, and Arm Circumferences, and Body Length of the Study Samples

Measurement	Male (n = 197)	Female (n = 205)	Total (n = 402)
Weight (g)	3079.1 ± 424.0	2941.9 ± 427.4	3009.0 ± 430.7
Chest (cm)	32.8 ± 1.8	32.2 ± 1.8	32.5 ± 1.8
Arm (cm)	10.7 ± 1.0	10.6 ± 1.0	10.6 ± 1.0
Length (cm)	49.5 ± 2.0	48.6 ± 1.9	49.0 ± 2.0

TABLE 2—Stepwise Regression for Birthweight Estimation

Step	Variable Entered	Coefficient	SE[a] of Coefficient	Error[b]	R^2 (%)[c]
1	Chest	207.2	5.4	199.8	78.5
	Constant	−3725.2			
2	Chest	149.6	8.3	183.3	82.0
	Arm	134.4	15.4		
	Constant	−3281.5			
3	Chest	104.8	9.0	167.0	85.1
	Arm	131.5	14.0		
	Length	57.4	6.3		
	Constant	−4607.0			

[a] SE = standard error.
[b] Standard error of estimate.
[c] R^2 = Coefficient of determination.

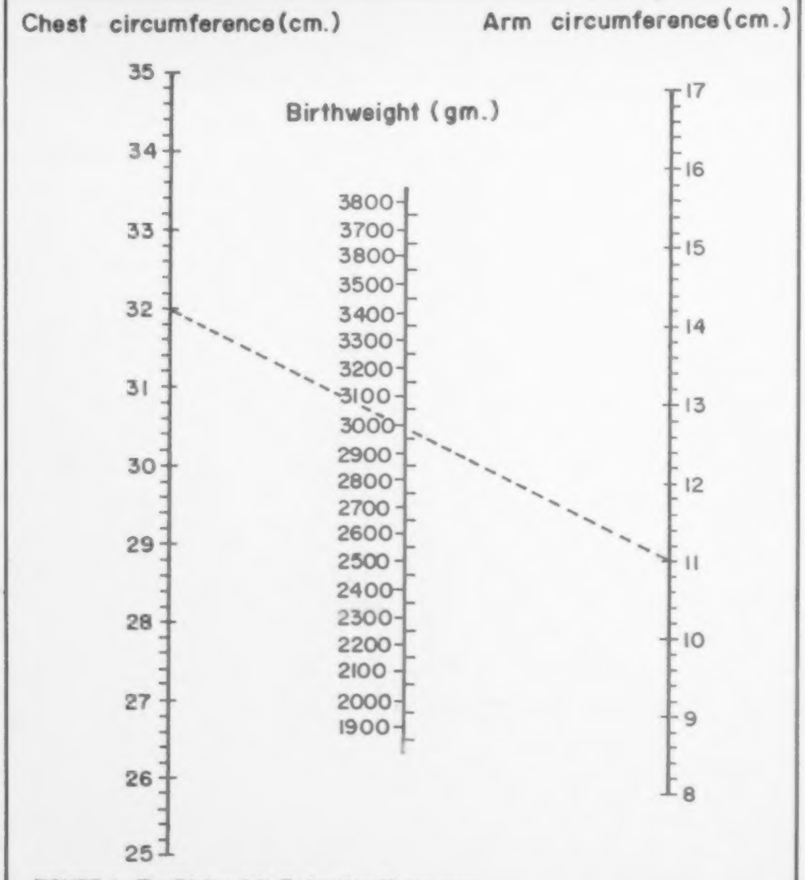

FIGURE 1—The Birthweight Estimation Nomogram

land was used for analysis. Measurements were taken within 24 hours of birth. The study population was derived from two sets of data. The first set consisted of 1612 newborn babies from the project "Low Birth Weight Newborn Babies of High Risk Mothers" at Chulalongkorn University Hospital, and at regional maternal and child health hospitals in Chiengmai, Nakornsawan, Rajburi, Yala, and Masharaj Hospital at Nakorn-srithamaraj. This was a case-control study, with a 1:2 ratio of low birthweight (LBW) to normal birthweight infants.

The second set of data consisted of 1610 newborn babies from Lerdsin Hospital and the Khon-Kaen University Hospital in Bangkok.

The first set was collected by nurses, using standardized measuring methods, and the second set was obtained by the doctors responsible for the project.

A diagnostic test was employed to reveal sensitivity and specificity of positive predictive values, negative predictive values, and accuracy of the new device in comparison with the gold standard (a baby metric beam balance) in screening for LBW infants.

Results

The data from the study of 402 newborn babies are summarized in Table 1. The correlation coefficients (r) of weight versus chest and arm circumferences and body length were 0.89, 0.82, and 0.78, respectively.

Although the combined chest and arm circumferences and length gave the best correlation with birthweight ($r = .92$), measuring the length of a newborn baby is usually impractical and may be subject to relatively large error. We therefore chose the second best equation (Table 2), using the combined chest and arm circumferences to estimate birthweight ($r = 0.90$, $r^2 = 82.0\%$, and standard error (SE) = ±183.3 g). Based on this equation, a nomogram was constructed (Figure 1) and, subsequently, a circular birthweight estimation device was developed (Figure 2).

The mean and standard deviations (SD) of chest and arm circumferences of 3222 newborn babies in various study centers are shown in Table 3. The mean of chest circumferences was about 31 cm, and arm circumferences was about 10 cm.

The differences of means between the weight obtained by standard scales and the estimated weight obtained by the new device (Figure 2) are also shown. The differences appear to be within ±187 g.

The diagnostic values of the new device, when the arbitrary point for picking LBW infants is set at 2500 g, are shown in Table 4. The sensitivity was lowest (56.9%) at Nakornsawan and highest (94.5%) at Chulalongkorn hospital. The specificity was lowest (86.1%) at Khon-Kaen and highest (98.6%) at Yala. The accuracy varied from 83.2% to 95.7%.

If the arbitrary point for defining LBW neonates is set at 2600 g on the new device, the sensitivity at all centers increased, ranging from 71.3% to 97.3%, the specificity ranged from 78.4% to 97.2%, and the accuracy ranged from 79.8% to 92.3%, as shown in Table 5. If the arbitrary point was set at 2700 g, the sensitivity increased to 100% at Chulalongkorn Hospital but the specificity declined to 58.7% at Khon-Kaen Hospital (Table 6).

These results showed that the optimum arbitrary point of all data combined for defining LBW neonates when using the new device should be 2600 g and can be confirmed by the receiver operating characteristic curve shown in Figure 3. The curve became steeper when the arbitrary point was changed from 2500 g to 2600 g, which means that the sensitivity increased markedly while only a small degree of specificity was lost. However, the curve became linear when the arbitrary point was 2700 g, which indicates a very small increase in sensitivity but marked loss of specificity.

Discussion

Birthweight of infants estimated from the Chulalongkorn birthweight estimation device differed by only ±187.4 g from weight obtained by a standard scale. This difference is not clinically significant unless the birthweights are below 2000 g.

The sensitivity and specificity of the new device varied when arbitrary points of LBW infants were set at 2500, 2600, and 2700 g. There was less variation in accuracy, however. The optimum arbitrary point on the new device, as determined by a receiver operating characteristic curve, was found to be 2600 g. The sensitivity increased to 100% when the arbitrary point was 2700 g, but the specificity was low, indicating that there would be an increased number of falsely identified LBW infants.

The Chulalongkorn birthweight estimation device appears promising as an appropriate substitute for obtaining the birthweight of newborns in places where scales are not readily available. While the theoretical error was ±183.3 g, the prac-

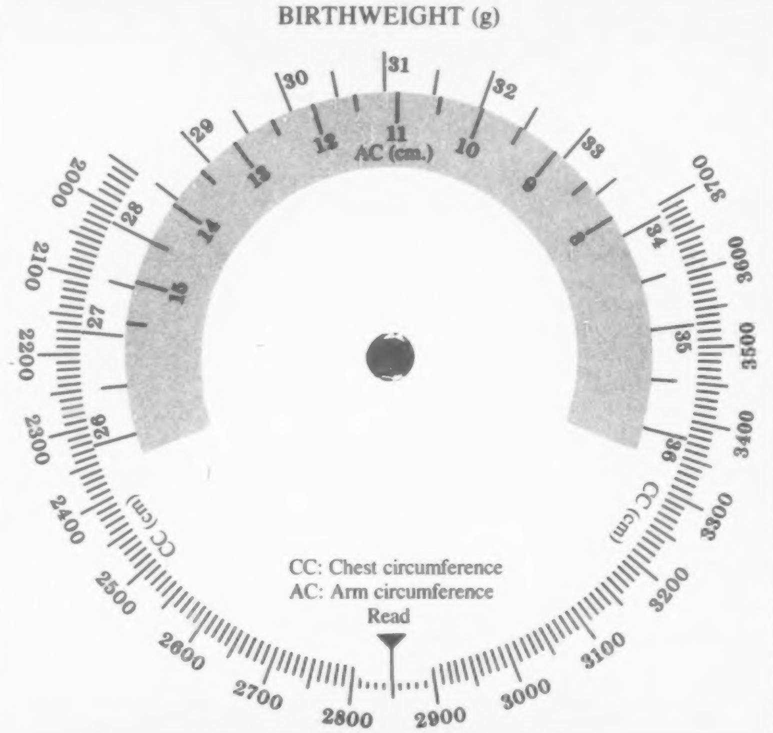

FIGURE 2—Chulalongkorn University Birthweight Estimation Device

TABLE 3—Mean ± Standard Deviation of Arm and Chest Circumferences and Weight of Newborn Babies

Location	n	Chest (cm)	Arm (cm)	Scale Weight	Estimated Weight	Difference of Mean
LS	606	32.3 ± 1.8	9.9 ± 0.9	2996.0 ± 378.7	2854.5 ± 331.4	141.5
KK	1004	31.6 ± 1.6	9.9 ± 0.7	2974.2 ± 360.8	2786.8 ± 321.0	187.4
CU	260	30.9 ± 2.2	9.9 ± 0.9	2828.6 ± 482.7	2668.6 ± 438.7	159.9
CM	339	31.3 ± 2.2	10.0 ± 1.0	2768.6 ± 482.6	2747.4 ± 440.5	21.2
NS	215	31.5 ± 2.3	10.6 ± 1.2	2887.5 ± 532.2	2861.4 ± 501.9	26.1
RB	215	31.4 ± 2.4	10.3 ± 1.2	2779.6 ± 554.8	2790.0 ± 501.5	8.7
YL	300	31.3 ± 2.4	10.8 ± 1.4	2759.2 ± 544.3	2857.1 ± 542.6	−97.9
NT	276	31.8 ± 2.5	10.7 ± 1.2	2797.6 ± 515.3	2925.2 ± 531.3	−127.6

Note. LS = Lerdsin General Hospital; KK = Khon-Kaen University Hospital; CU = Chulalongkorn University Hospital; CM = Chiengmai MCH Hospital; NS = Nakornsawan MCH Hospital; RB = Ratchaburi MCH Hospital; YL = Yala MCH Hospital; NT = Nakorn-srithamaraj MCH Hospital.

TABLE 4—Diagnostic Values for Detecting Low Birthweight Babies (<2500 g) When the Arbitrary Point Is Set at 2500 g on the New Device

Diagnostic Values	LS	KK	CU	CM	NS	RB	YL	NT
Sensitivity	79.4	88.1	94.5	75.5	56.9	65.3	65.8	57.4
Specificity (%)	97.6	86.1	93.0	96.1	96.5	97.5	98.6	98.4
Accuracy (%)	95.7	86.2	93.5	89.4	83.2	86.7	87.4	84.4
+PV (%)	79.4	37.0	—	—	—	—	—	—
−PV (%)	97.6	98.8	—	—	—	—	—	—

Note: +PV, positive predictive value; −PV, negative predictive value; Positive and negative predictive values of case-control study cannot be calculated, since estimates of prevalence of low birthweight were not available.

LS = Lerdsin General Hospital; KK = Khon-Kaen University Hospital; CU = Chulalongkorn University Hospital; CM = Chiengmai MCH Hospital; NS = Nakornsawan MCH Hospital; RB = Ratchaburi MCH Hospital; YL = Yala MCH Hospital; NT = Nakorn-srithamaraj MCH Hospital.

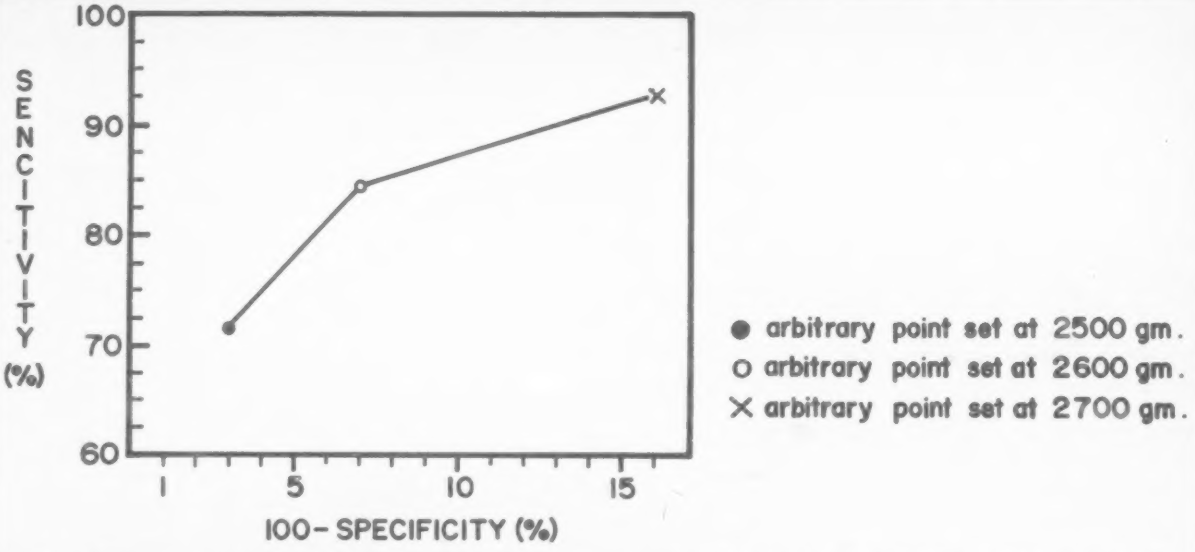

FIGURE 3—Receiver Operating Characteristics Curve of Birthweight Determination in Diagnosing Low Birthweight (All Data Combined, N = 3222)

TABLE 5—Diagnostic Values for Detecting LBW Babies (<2500 g) When the Arbitrary Point is Set at 2600 g on the New Device

Diagnostic Values	LS	KK	CU	CM	NS	RB	YL	NT
Sensitivity (%)	87.3	95.2	97.3	89.5	80.6	76.2	81.6	71.3
Specificity (%)	92.9	78.4	85.6	93.8	93.7	96.5	95.7	97.2
Accuracy (%)	92.3	79.8	88.9	92.3	89.3	89.0	91.0	88.4
+PV (%)	59.1	—	—	—	—	—	—	—
−PV (%)	98.4	—	—	—	—	—	—	—

Note. +PV, positive predictive value; −PV, negative predictive value; Positive and negative predictive values of case-control study cannot be calculated, since estimates of prevalence of low birthweight were not available.

LS = Lerdsin General Hospital; KK = Khon-Kaen University Hospital; CU = Chulalongkorn University Hospital; CM = Chiengmai MCH Hospital; NS = Nakornsawan MCH Hospital; RB = Ratchaburi MCH Hospital; YL = Yala MCH Hospital; NT = Nakorn-srithamaraj MCH Hospital.

TABLE 6—Diagnostic Values for Detection of LBW Babies (<2500 g) When the Arbitrary Point is 2700 g on the New Device.

Diagnostic Values	LS	KK	CU	CM
Sensitivity (%)	92.1	96.6	100.0	93.9
Specificity (%)	80.5	58.7	74.3	81.8
Accuracy (%)	81.7	62.0	81.5	85.8
+PV (%)	35.6	—	—	—
−PV (%)	98.9	—	—	—

Note. +PV, positive predictive value; −PV, negative predictive value; positive and negative predictive values of case-control study cannot be calculated, since estimates of prevalence of low birthweight were not available.

LS = Lerdsin General Hospital; KK = Khon-Kaen University Hospital; CU = Chulalongkorn University Hospital; CM = Chiengmai MCH Hospital.

tical error was ±187.4 g when compared with standard weighing scales.

The diagnostic test revealed the value of the new device. The optimum arbitrary point for screening an LBW infant was 2600 g, at which the sensitivity was 97.3%, the specificity was 97.2%, and the accuracy was 92.3%. This new device is inexpensive and convenient and is a good example of appropriate technology for primary health care. Use of the device should be encouraged throughout developing countries such as Thailand. ☐

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Oral Rehydration Program Evaluation by Quality Assurance Sampling in Rural Haiti

June N. Pierre-Louis, MPH

ABSTRACT

An oral rehydration program in rural Haiti was evaluated by quality assurance sampling. The quality assurance sampling method and its application are described. The indicators measured were knowledge of the oral rehydration salts packet, knowledge of preparation of the solution from the packet, oral rehydration therapy use, and knowledge of preparation of the solution from salt and sugar. Coverages of the first two indicators were adequate, coverages of the latter two were inadequate. The method is a useful low-cost approach to evaluation of program coverage. (*Am J Public Health.* 1991;81:1205–1207)

Introduction

Oral rehydration therapy (ORT) is a cost-effective treatment for diarrheal disease.[1] While the scientific basis for oral rehydration is established, and programs are in place, many programs have not been evaluated at the local level.

In rural Haiti local programs typically target populations of less than 50,000.[2] To meet the need for a low-cost, population-based survey method, quality assurance sampling (QAS), a small sample method,[3] was used to evaluate a local ORT program in Haiti. The QAS method and its application in Haiti are described.

Methods

Study Population

Freres is a rural area located 20 km from Port-au-Prince, the capital of Haiti. The population (13,750 in 1987) lives on dirt roads within one km of a partly paved main road which bisects the area. Freres is targeted by a primary health care program managed by a private hospital in the area. The program maintains a census of the target population.

Promotion of ORT takes place at community meetings and on the radio. Health workers teach mothers to prepare the solution of oral rehydration salts (ORS) from a packet, and from salt and sugar, and to use ORT when their children under three years of age have diarrhea.[2]

Indicators

Widely accepted indicators of program outcome were selected by managers and funders of oral rehydration programs in Haiti and included: ORT use, knowledge of the packet, knowledge of preparation of the solution from the packet, and from salt and sugar.[4,5] The use rate was defined as the proportion of episodes of diarrhea in children under three years of age receiving ORT, whether the solution was prepared from the packet or salt and sugar.[4] The rate of knowledge of preparation of the solution was defined as the proportion of mothers of children under age three knowing the correct quantities of ingredients in a solution prepared from the packet or from salt and sugar.

The survey questionnaire is given in the Appendix.

Quality Assurance Sampling Method

The QAS method was first used for quality control in manufacturing. Product quality was assessed based on the probability of observing a particular proportion of defective items in a sample relative to a null hypothesis stating a proportion of defective products. This probability is determined by the binomial distribution. The goal of the QAS method is hypothesis testing rather than parameter estimation. The major advantage of QAS is the sample size which is small.

Address reprint requests to June N. Pierre-Louis, MPH, Division of Nutritional Sciences, Cornell University, Ithaca, NY 14853. The author was affiliated with Institut Haitien de l'Enfance, Delmas, Haiti. This paper, submitted to the Journal December 1, 1989, was revised and accepted for publication January 23, 1991.

TABLE 1—Hypothesized Proportions of Nonrecipients and Recipients, Null Hypotheses and Sample Sizes, Oral Rehydration Program Evaluation, Quality Assurance Sampling, Freres, Haiti, 1987

Indicator	Hypothesized Proportion of Nonrecipients P_0	Hypothesized Proportion of Recipients/ Short-term Program Goal $1-P_0$	Null Hypothesis	Sample Size n	Critical Number Nonrecipients d^*
ORT Use	0.95	0.05	$P_0 \geq 0.95$	20	17
Knowledge of Packet	0.30	0.70	$P_0 \geq 0.30$	45	9
Knowledge of Preparation from Packet	0.70	0.30	$P_0 \geq 0.70$	45	27
Knowledge of Preparation from Salt and Sugar	0.90	0.10	$P_0 \geq 0.90$	45	37

TABLE 2—Coverage and Estimated Proportions of Nonrecipients and Recipients, Oral Rehydration Program Evaluation, Quality Assurance Sampling, Freres, Haiti, 1987

Indicator	Sample Number Nonrecipients d	Critical Number Nonrecipients d^*	Coverage	Estimated Proportion of Nonrecipients	Estimated Proportion of Recipients
ORT Use	19	17	Inadequate	≥0.95	<0.05
Knowledge of Packet	8	9	Adequate	<0.30	≥0.70
Knowledge of Preparation of the Solution from Packet	25	27	Adequate	<0.70	≥0.30
Knowledge of Preparation of the Solution from Salt and Sugar	40	37	Inadequate	≥0.90	<0.10

The rate or coverage of use of ORT in the target population, for instance, is determined to be adequate or inadequate based on finding a certain proportion of children under age three not receiving the benefits of ORT in a sample. The null hypothesis is that ORT coverage is inadequate; that is, the proportion of nonrecipients in the sample is greater than some hypothesized proportion of nonrecipients based on short-term program objectives (see Table 1). The null hypothesis is so stated to control the probability of concluding that a community has adequate health coverage when it is actually inadequate. The probability of such a false negative result was set at 0.10, (α). This error is more serious than a false positive error (β) in which the conclusion is that health coverage is inadequate when it is really adequate.

The following expression holds:[3]

$$\alpha = \Pr\{d \leq d^*\}$$

where d is the observed number of nonrecipients in the sample and d^* is the critical number of nonrecipients in a given sample size. When $d \leq d^*$ the null hypothesis that coverage is inadequate is rejected and it is concluded that coverage is adequate.

The critical number of nonrecipients in the sample was calculated by:[3]

$$d^* = nP_0 - z_\alpha\,[nP_0(1 - P_0)]^{1/2}$$

where n is the sample size; P_0 is the hypothesized proportion of nonrecipients; $1 - P_0$ is the hypothesized proportion of recipients; and z equals 1.28 for α equal to 0.10 using a one-sided test. Fractions were rounded down to the next whole number.

Null hypotheses and values of P_0, $1 - P_0$, n and d^* are given in Table 1 for each indicator.

Sample sizes were obtained from tables developed by Lemeshow, *et al.*,[6] of (n, d^*) pairs for selected values of α, β, and P_0. Values of β are estimates of the difference between true and hypothesized proportions of nonrecipients. A sample size of 45 was chosen for the indicators except ORT use. A mother was asked about ORT use only if she said her child had diarrhea in the past two weeks to minimize recall bias. Sixty-three mothers were interviewed, of which 20 had children under age three who had diarrhea in the past two weeks.

A systematic sample was drawn from the census of the target population maintained by the primary health care program.

The nonresponse rate was 9 percent.

Results

Coverages of ORT use and of knowledge of preparation of the solution from salt and sugar of mothers of children less than three years of age were determined to be inadequate (Table 2). The ORT use rate was estimated to be <0.05. The rate of knowledge of preparation of the solution from salt and sugar was estimated to be <0.10.

Coverages of knowledge of the packet and of preparation of the solution from the packet of mothers of children under age three were determined to be adequate. The rate of knowledge of the packet was estimated to be ≥0.70. The rate of knowledge of preparation of the solution from the packet was estimated to be ≥0.30.

Discussion

The QAS survey in Freres revealed that the primary health program had inadequate coverage of ORT use in the target population despite adequate rates of knowledge of the packet and of preparation of the solution from the packet. The probability that the latter two rates are false negative results is ≤0.10.

The estimated ORT use rate in Freres in 1987 (<0.05) was much lower than the rate for Haiti in 1985 (0.23 of mothers of children less than five).[7] The estimated rate of knowledge of the packet in Freres

(≥0.70) was similar to the national rate in 1985 (0.58).

The QAS evaluation was relatively low-cost because the sample size was small. One interviewer was required for the survey and the evaluation was completed in two days. The QAS method is useful for evaluation of local oral rehydration programs in Haiti. ☐

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APPENDIX

Survey Questionnaire, Quality Assurance Sampling Survey, Freres, Haiti, 1987.

1. Did (name of the child) have diarrhea in the past two weeks? (If not, pass to question 3).
Yes No

2. What did you give (name of the child)?
(Circle all spontaneous responses of the mother. Do not mention alternatives.)
Home Remedy Doctor's Medicine Packet Solution
Salt and Sugar Solution Nothing

3. Do you recognize this packet? (Show the ORS packet.)
What is it called? ____________

4. How do you prepare the packet ORS solution? (Circle all spontaneous responses of the mother. Do not mention alternatives.)
3 Kola Bottles of Water 1 Rum Bottle of Water 1 ORS Packet
Solution Boiled Other ____________ Doesn't Know

5. How do you prepare the salt and sugar ORS solution? (Circle all spontaneous responses of the mother. Do not mention alternatives.)
2 Large Spoons of Sugar 1 Small Spoon of Salt
1 Rum Bottle of Water 3 Kola Bottles of Water
2 Small Spoons of Sugar 1 Pinch of Salt
1 Kola Bottle of Water Solution Boiled
Other ____________ Doesn't Know

Impaired and Disabled Elderly in the Community

Amasa B. Ford, MD, Ann W. Roy, PhD, Marie R. Haug, PhD, Steven J. Folmar, PhD, and Paul K. Jones, PhD

ABSTRACT

A nine-year representative, longitudinal study of 1,598 urban elderly shows that two to eight times as many impaired or disabled are cared for in the community as in institutions. Younger age, male gender, better income, and living with others, especially children, favor continuing care in the home. (*Am J Public Health.* 1991; 81:1207–1209)

Introduction

The purpose of the present study is to report information about the distribution between community and long-term institutional care of a representative sample of urban elderly over a nine-year period. We identify supporting factors that have enabled from two to eight times as many of the most impaired and disabled to remain in the community as were admitted to long-term nursing home care.

Methods

In 1975 a random sample of 1,598 noninstitutionalized persons age 65 years and older, living in the City of Cleveland, Ohio, were interviewed in their homes using the Duke OARS instrument.[1,2]

Nine years later, in 1984, the vital status of all but 13 individuals (0.8 percent) was ascertained. Among the 805 who were known to have survived, 647 (80.4 percent) were reinterviewed, using the same instrument. Information sufficient to determine institutionalization was ob-

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From the Department of Epidemiology and Biostatistics, School of Medicine, and the Center on Aging and Health, Case Western Reserve University; and the Carolina Population Center, University of North Carolina, Chapel Hill.

Address reprint requests to Amasa B. Ford, MD, Department of Epidemiology and Biostatistics, School of Medicine, Case Western University, Cleveland, OH 44106. This paper, submitted to the Journal May 29, 1990, was revised and accepted for publication January 24, 1991.

TABLE 1—Survivors (n = 589) Remaining in the Community or Institutionalized Long-term, 1975–84, According to Health Impairment and Disability in 1984

	Remaining in Community (n)	Institutionalized Long-term (n)	Percent of Impaired or Disabled Remaining in Community
Survivors	526	63	
Physical health			
4+ chronic conditions	124	16	88.6
<4 chronic conditions	402	47	
Cognitive impairment			
4+ errors	63	26	70.8[a]
<4 errors	443	10	
Psychological distress			
7+ symptoms	95	42	69.3[a]
<7 symptoms	431	21	
Functional disability			
dependent 1+ ADL	111	60	64.9[a]
not ADL dependent	415	3	

[a] $p < .001$ by chi square

TABLE 2—Logistic Regression of Demographic and Independent Variables, Measured in 1975, on Dependent Variable: Remaining in Community 1975–84, for Survivors (n = 589) and Deceased (n = 632)[a]

Variables	Odds Ratio	95% CL
Age (continuous, by year)	.93	.90–.96
Gender (M = 1)	2.70	1.72–4.16
Race (B = 1)	1.38	.87–2.17
Education (>8 yr. = 1)	.73	.50–1.07
Income ($5000+ = 1)	2.28	1.46–3.54
Lives with child(ren) only (= 1)	3.45	1.85–6.43
Physical health impairment (4+ chronic conditions = 1)	1.11	.68–1.82
Cognitive impairment (4+ errors = 1)	.85	.46–1.56
Psychological distress (>7 of 15 symptoms)	1.53	.87–2.70
Functional disability (help with 1+ ADL = 1)	.79	.52–1.20
Died between 1975 and 1984 (= 1)	.56	.38–.83

[a] 82 cases omitted from analysis because of missing data.

tained for 589 of the living respondents (91 percent). By 1984, 780 individuals were known to have died. Telephone interviews with next of kin or other informants were obtained for 639 deceased, with data about institutionalization secured for 632 (81 percent).

Demographic information was collected according to the OARS instrument. Impaired physical health was indicated by the presence of four or more self-reported chronic conditions from a list of the 12 most prevalent conditions.[3] Impaired cognitive function was measured by a score of four or more errors on a 10-item test of cognitive function.[4] "Psychological distress" was measured as seven or more symptoms of a list of 15 on a Short Psychiatric Evaluation Schedule.[5] Functional disability was measured by need for help with one or more of six activities of daily living: walking, going outside, bathing, dressing, transfer (i.e. getting in or out of bed or chair), or eating.[6]

The definition of long-term institutionalization was one or more institutional admissions totaling 30 days or more during the nine-year observation period, or prior to death. The rest of the subjects were characterized as "remaining in the community," even though they may have experienced short-term admissions. The object was to distinguish those who continued to receive care mainly outside of institutions from those who entered long-term or indefinite institutional care.

Results

Comparison of the 1984 survivors with figures extrapolated from the 1980 census of Cleveland shows that the subjects, now aged 74 years and older, remained representative of the city population aged 75 and over.[7]

By 1984, 76 of the 647 responding survivors (11.8 percent) were living in an institution at the time of the interview, and a total of 63 (10.7 percent) of all respondents, including institutionalized and noninstitutionalized, had been admitted to a long-term care institution for 30 days or more. Of 731 persons who died during the nine-year observation period and for whom death certificates were available, 28.3 percent had a nursing home as their usual place of residence at the time of death.

Considering first the 589 survivors with complete data and their state of health impairment and functional disability at the time of reinterview in 1984, it appears that mental health impairment and functional disability were associated with institutionalization, but that the chances of remaining in the community were not affected by the presence of multiple chronic conditions. Expressed as a ratio of numbers of community residents to numbers of institutional residents, the figures are 7.8 for physical health impairment, 2.4 for cognitive impairment, 2.3 for psychological distress, and 1.8 for functional disability (Table 1).

The next stage of analysis was to ask whether it would have been possible to predict at the beginning of the study who would experience long-term institutionalization and who would remain in the community. A multivariate logistic regression was performed, using remaining in the community as the dependent variable, with independent variables as indicated in Table 2.

Factors significantly favoring continued care in the community include: younger age, male gender, income of greater than $5,000 (1975), and living with child(ren) only. Since previous analyses had pointed to living with others as favoring community care, three separate regressions were carried out, using different definitions of this variable. The odds ratio resulting were: lives with spouse, or children and spouse: 1.76 (95% CL = 1.08, 2.89), lives with others: 2.23 (95% CL = 1.54, 3.22), and lives with child(ren) only: 3.45 (95% CL = 1.85, 6.43).

Somewhat surprisingly, none of the four measures of impairment or disability served to distinguish between those who remained in the community and those who experienced long-term institutionalization. Those who died during the study period were less likely to have avoided long-term institutionalization.

Discussion

These results indicate that socioeconomic factors predominate as predictors of long-term care in the community. The way these factors operate is fairly clear. Relative youth implies better resources and coping skills. Males are more likely to have a surviving spouse. Better income allows for better housing and purchase of services, and children can be both care providers and income producers. In other recent longitudinal studies age, female gender, and living alone or absence of spouse have been shown to predict institutionalization.[8,9] Notably in the present study, when elders live with them, children appear to enhance the possibility of remaining in the community more than do living spouses or nonspecified "others."

Although, in this study, impairment and disability do not appear to predict whether community or institutional care will be used, it must be remembered that some part of nine years intervened between the measurement of dependent and independent variables. Observations closer to the decision-point might have shown that mental health impairment and functional disability were more difficult to care for at home than impairment of physical health alone, as suggested by the data in Table 1. ☐

Acknowledgments

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This report was presented in preliminary form at the World Congress of Gerontology, Acapulco, Mexico, June 22, 1989.

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Recent Trends in the Incidence of Toxic Shock Syndrome in Northern California

ABSTRACT

The incidence of toxic shock syndrome in women members of a large prepaid medical care program in Northern California was 1.5 cases per 100,000 in a period after removal of tampons containing polyacrylate rayon and reductions in tampon absorbency. This rate was lower, but not significantly lower, than the rate of 2.2 per 100,000 in the prior interval. It was higher, but not significantly higher, than the rate of 0.4 per 100,000 in the era before "superabsorbent" materials were introduced into tampons. The incidence in men has been stable at about 0.1 cases per 100,000 for the 15-year period from 1972 though 1987. (*Am J Public Health.* 1991;81:1209–1211)

Diana B. Petitti, MD, MPH, and Arthur L. Reingold, MD

Introduction

We previously reported the incidence of toxic shock syndrome (TSS) through 1985 in members of the Northern California Kaiser Permanente Medical Care Program ages 15–34 years.[1–3] Here, we present a final update on TSS in this population, presenting data for 1986 and 1987. The incidence of TSS in these latter two years is of particular interest, because polyacrylate rayon, a material included in tampons to increase their absorbency, was removed from tampons marketed in the United States in March 1985 and because there were further declines in the absorbency of tampons overall. Since high absorbency has been linked with an increase risk of TSS in tampon users,[4] one would expect the incidence of TSS to decline in women after reductions in the absorbency of tampons brought about, at least in part, by removal of polyacrylate rayon from tampons.

Methods

Our method of case ascertainment, in which medical records were reviewed to identify recognized and unrecognized cases of TSS, was described in detail in a

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From the Division of Research, Kaiser Permanente Medical Care Program, Oakland, CA (Petitti) and the Department of Biomedical and Environmental Health Sciences, School of Public Health, University of California, Berkeley (Reingold).

Address reprint requests to Diana B. Petitti, MD, MPH, Department of Family and Community Medicine, AC-9, Box 0900, University of California at San Francisco, San Francisco, CA 94143. This paper, submitted to the Journal May 25, 1990, was revised and accepted for publication January 8, 1991.

TABLE 1—Incidence of Definite TSS in Females Ages 15–34 in Relation to Certain Milestones in Tampon Formulation and Marketing

Time Periods	Description	Number of Cases	Incidence* (95% Confidence Interval)
1. 1972–June 1977	Prior to introduction of superabsorbent materials	5	0.4 (0.1,1.0)
2. July 1977–April 1979	After superabsorbents introduced, but before Rely marketing in California	8	1.5 (0.6,2.9)
3. May 1979–September 1980	Rely and other superabsorbents in California	11	2.4 (1.2,4.2)
4. October 1980–March 1985	After Rely removal from the market but before removal of polyacrylate rayon	34	2.2 (1.5,3.0)
5. April 1985–December 1987	After removal of polyacrylate rayon products and reductions in absorbency	14	1.5 (0.8,2.6)

* per 100,000 women-years

prior publication.[1] Briefly, we used computer-stored information from records of discharges from Northern California Kaiser Permanente Medical Care Program facilities for the period 1972 through 1987 to identify men and women ages 15–34 years hospitalized with illnesses that might be confused with TSS, or, for later years, with diagnosed TSS. Records for qualifying hospitalizations were reviewed by a single trained medical record technician, who determined whether or not the illness met the case definition for TSS developed by the Centers for Disease Control (CDC), except that desquamation and orthostatic dizziness were not considered criteria because desquamation was not reliably recorded in the medical records and because orthostatic hypotension was not an official criterion for TSS when the study began. Records for all illnesses that might represent TSS were then reviewed independently by each of the co-authors, who classified them as definite TSS, probable TSS, or not TSS. Definite TSS cases were illnesses that met all the CDC criteria as TSS, except desquamation. Probable cases were illnesses that had rash, fever, and hypotension, but involved only two organ systems.

The denominators for the calculation of the incidence rates were person-years of membership in the Northern California Kaiser-Permanente Medical Care Program. Ninety-five percent confidence intervals were estimated based on the assumption that the incidence of TSS is distributed as a Poisson. Statistical tests of the hypotheses about rates of TSS were carried out based on the comparison of two Poisson variables.

Results

Table 1 shows the incidence of definite TSS in women in relation to certain milestones in tampon formation and marketing. The incidence of TSS was statistically significantly higher (p = 0.0003) after introduction of superabsorbent materials and Rely tampons into the California market (period 3) than in the period prior to these two changes in tampon for-

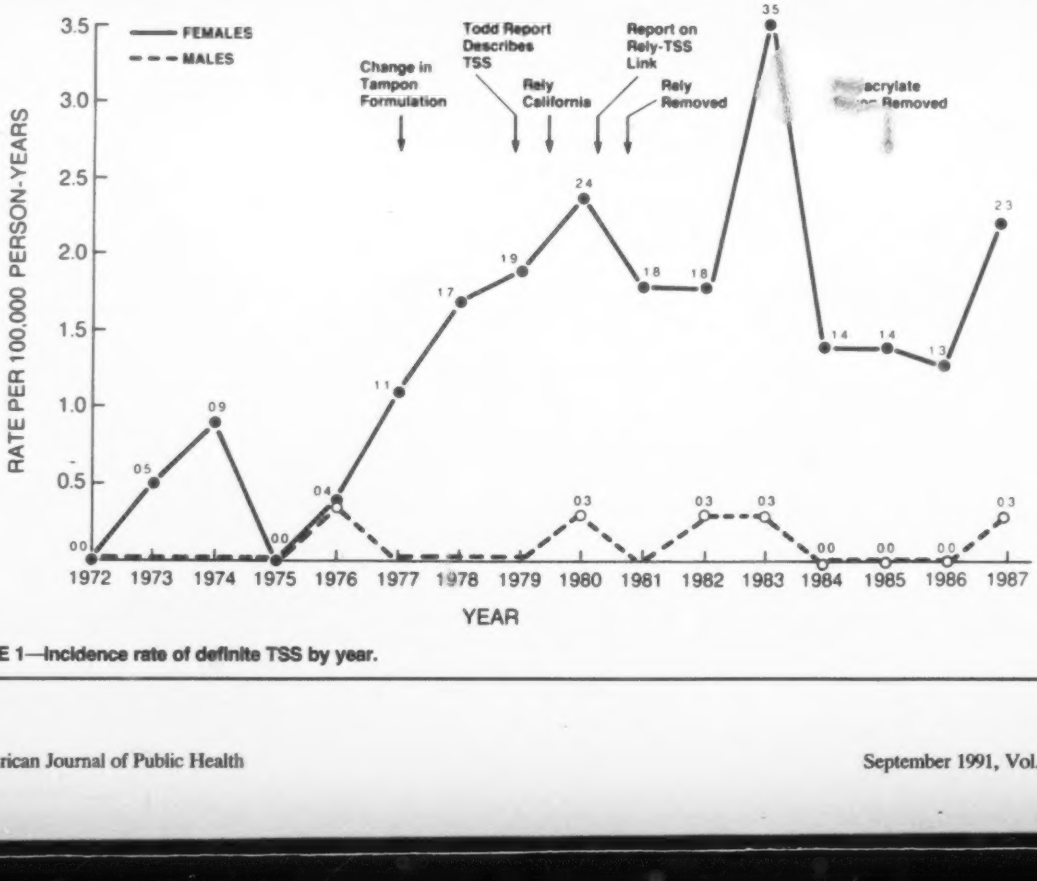

FIGURE 1—Incidence rate of definite TSS by year.

mulation and marketing (period 1). The incidence of definite TSS in women after removal of Rely and polyacrylate rayon from remaining tampons (period 5) was lower, but not significantly lower ($p = 0.13$) than in the prior two intervals (periods 3 and 4). The incidence rate after removal of polyacrylate (period 5) was also not statistically significantly higher ($p = 0.13$) than in the period prior to introduction of superabsorbent materials (period 1).

Information on menstrual status at the onset of symptoms was retrievable from the medical records of all 12 definite TSS cases in women in 1986 and 1987 (Figure 1). Eight of these cases (67 percent) were menstrual (95% CI = 49, 85). This percentage is the same as it was in prior years in this population.[2]

The incidence of TSS in men has been stable at about 0.1 cases per 100,000 throughout the 15-year period covered by this surveillance project.

Discussion

The incidence of TSS began increasing in about 1977 in temporal relation to changes in tampon formulation that made them more absorbent. The rarity of TSS in this population limits the ability to draw firm conclusions about the incidence of TSS in relation to the removal of Rely tampons from the market, the subsequent removal of polyacrylate rayon, and the reductions in the absorbency of tampons that occurred simultaneously. The data are suggestive of a decline in incidence in relation to removal of polyacrylate rayon and accompanying reductions in absorbency. However, these changes in tampon formulation have clearly not eliminated TSS in women.

It is reasonable to conclude TSS is very rare in men. It also appears that in this population the percentage of menstrual cases is stable at about two-thirds of all cases in women. The incidence in both men and women in this population is very low, and TSS is not an important public health problem at this time. ☐

Acknowledgments

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Hospital Charges in Five Washington Counties Jump Average of Nearly 37%

Hospital charges in the state of Washington's five largest counties increased by an average of nearly 37% between 1987 and 1989, according to a report by the Washington State Department of Health.

The report, "Inflation Measures in Hospital Charges," covers 15 separate categories, from coronary bypass surgery to Caesarean delivery, at hospitals in King, Pierce, Snohomish, Thurston, and Spokane Counties.

"Hospital charges climbed at a significant rate over that 3-year period," said State Health Secretary Kristine M. Gebbie. "This report confirms that the issues surrounding the inflation of hospital charges are complex and that the upper range of these inflation measures are well above the consumer price index or even average increases in medical charges."

Increases in charges over the 3-year period range from 9.9% for treating alcohol and drug dependency to 67.7% for newborn children with major medical problems. Annual inflation figures for those same categories of treatment were 4.1% and 32.8%, respectively, with a median annual inflation rate of 10.5% covering all 15 categories.

The overall average increase for the 15 categories of services was 36.6% over the 3-year period.

The amount of inflation varied by category of service and the county. For example, average charges for cardiac bypass surgery in King County jumped from about $17,500 in 1987 to just over $22,200 in 1989, whereas in Pierce County the charges for similar procedures increased from just under $23,000 to more than $30,000. In Spokane County, similar bypass surgery jumped from about $18,800 to about $21,500.

For this type of bypass surgery, the overall average charge increased from $19,038 to $23,482, or about 23.3%. Annually, this meant the inflation rate for this procedure was 10.7%. The greatest total increase over the 3-year period was 30.8% in Pierce County, and the smallest gain was in Spokane County with 14.6% over 3 years.

Another category, care for newborn children without significant problems, represents the greatest single area of competition between hospitals, according to the state agency's report.

The average charge for newborn care within the five counties increased from $478 for the first half of 1987 to $555 for the second half of 1989, or about 16%. This level of increase did not occur in all counties. Snohomish and Spokane counties posted 20.8% and 28.5%, respectively.

Two measures of inflation are used extensively throughout the report: total change in the average charge from the first half of 1987 through the last half of 1989; and the average annual rate for 1988 and 1989. The report also measures the average level of inflation in charges covering all patients and all categories.

In addition to bypass surgery and Caesarean section deliveries, the 15 categories of service include chemotherapy, reattaching major joints or limbs, and psychiatric care, including the treatment of schizophrenia and manic depression.

Anyone who wants a copy of the report may contact the State Department of Health, Office of Hospital Data Systems, Mail Stop FJ-21, Olympia, WA, 98504, or phone 206/753-1990 during regular business hours.

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encompassing approach, a renewal of enthusiasm for control of the "other STDs." ☐

Richard B. Rothenberg, MD

Richard B. Rothenberg is with the Center for Chronic Disease Prevention and Health Promotion, Centers for Disease Control, Rhodes Building, Koger Center, 3005 Chamblee-Tucker Road, Atlanta, GA 30341. Requests for reprints should be sent to him at the above address.

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Erratum

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The reference information for this article printed at the bottom of each page (running feet) was incorrect; it should have read August 1991, Vol. 81, No. 8.

Sociodemographic Distribution of Gonorrhea Incidence: Implications for Prevention and Behavioral Research

Roselyn J. Rice, MD, Pacita L. Roberts, MS, H. Hunter Handsfield, MD, and King K. Holmes, MD, PhD

ABSTRACT

Background. Despite a declining incidence during the AIDS era, gonorrhea remains the most frequently reported communicable disease in the United States.

Methods. During 1986 and 1987 we supplemented gonorrhea case reporting with laboratory surveillance in King County, Washington. Incidence rates were correlated with demographic variables.

Results. Overall incidence of gonorrhea was similar for men and women, but highest for 16- to 21-year-old females and urban Seattle residents. Incidence rates by ethnicity were Blacks, 3033; Native Americans, 843; Hispanics, 617; Asians, 190; and Whites, 121. Census tracts representing the lowest socioeconomic status (SES) quartile accounted for 58% of reported gonorrhea. Black female teenagers residing in the lowest SES urban areas had highest incidence rates: aged 14 to 15, 3.4%; 16 to 17, 10.4%; 18, 17.0%; and 19, 15.4%. Rates in female teenagers were even higher after adjustment for estimated proportion of those who were sexually experienced.

Conclusions. Gonorrhea incidence is associated with age, gender, ethnicity, SES, and residence. Identification of populations at highest risk for gonorrhea can direct interventions against all sexually transmitted diseases. Clearly, interventions to alter high-risk behaviors must be initiated in early adolescence. (*Am J Public Health*. 1991;81:1252–1258)

Introduction

With the advent of effective contraception, the onset of sexual maturity in the baby boom generation, and changes in sexual behavior, reported cases of gonorrhea increased at about 15% per year from the 1960s through the mid-1970s in the United States.[1,2] Reported cases have gradually fallen since 1975.[3] After 1982 a steady decline in rectal gonorrhea in men was observed in King County, Washington, and in other areas of the United States,[4,5] implying a decline in gonorrhea among homosexually active men. Further analysis of trends in occurrence of gonorrhea by age, gender, and race for the period 1980 through 1988 in King County has shown more than a threefold decline in cases occurring in White men, little change in White women, but a steady increase in Black men and women from 1982 through 1988 (authors' unpublished data). Among Blacks, the greatest increase in annual cases has occurred among teenagers. This unexpected increase in gonorrhea in Blacks and the concurrent leveling off of reported cases in White women after nearly a decade of decline led us to undertake a 2-year intensive surveillance of gonorrhea in Seattle-King County to define incidence rates in population subgroups according to age, gender, race or ethnicity, residence (urban, suburban, or rural), and socioeconomic status (SES). The results are useful not only for defining more precisely a current profile of individuals at highest risk for gonorrhea, but also as a proxy measure of the distribution of sexual behaviors that place individuals at risk for all sexually transmitted diseases, including the acquired immunodeficiency syndrome (AIDS).

Methods

Surveillance and Case Reporting

King County, Washington (population 1.4 million, 2130 square miles), encompasses the city of Seattle (population 496 000). Health care providers in King County are required to report all gonorrhea cases; compliance is 100% by Seattle-King County Department of Public Health (DPH) clinics, but variable by other clinics and the private health care sector.

Surveillance

Beginning in 1985, presentations to microbiology laboratory directors, monthly bulletins of the DPH, newsletters and bulletins of the King County Medical Society, and hospital staff meetings were used to publicize the enhanced gonorrhea surveillance project and encourage compliance with effective culture diagnosis and case reporting. Routine case reporting was supplemented by active laboratory surveillance for *Neisseria gonorrhoeae* in

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All authors are with the Department of Medicine, Harborview Medical Center, and the University of Washington School of Medicine; H. Hunter Handsfield is with the Seattle-King County Department of Public Health, Seattle; and Roselyn Rice is with the Division of Sexually Transmitted Diseases Laboratory Research, Center for Infectious Diseases, Centers for Disease Control, Atlanta

Requests for reprints should be sent to King K. Holmes, MD, PhD, Harborview Medical Center, 325 Ninth Avenue, Seattle, WA 98104.

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laboratories in King County through a system established with the DPH central laboratory. Hospital, private physician, and commercial laboratories were asked to submit all positive and confirmatory culture isolates for *N. gonorrhoeae* to the DPH laboratory for confirmation and antimicrobial susceptibility screening tests.

The DPH processed 56 353 cultures for *N. gonorrhoeae* in 1986 and 52 844 in 1987. During 1986 and 1987, gonococcal isolates were received from 14 hospitals, 4 private microbiology laboratories, 2 student health clinics, 2 military health clinics, all 12 DPH clinics, 22 community clinics (partially supported by DPH), the Seattle-King County Jail and Juvenile Detention Center, 6 private family planning clinics, and numerous physicians' offices. Of the cultures confirmed as positive for *N. gonorrhoeae*, 47.0% were submitted from DPH clinics, 22.2% from private physicians, 20.4% from hospitals, 6.0% from community clinics, 3.6% from private family planning clinics, 0.6% from student health clinics, and 0.1% from military clinics. Matching of positive cultures confirmed by the DPH laboratory with submitted case reports from all public and private providers indicated that cultures were submitted to the DPH laboratory for confirmation from 78% of the reported gonorrhea cases in the study population.

Case Reporting

Gonorrhea case reports were submitted to the DPH using Centers for Disease Control (CDC) form 73.688 (revised 1/86). Following notification of a case of gonorrhea by case report form or as a result of laboratory surveillance, when necessary, DPH staff contacted health care providers for additional demographic information on age, gender, race or ethnicity, address, anatomic site of infection, date of infection, date of treatment, and clinical diagnosis (asymptomatic gonorrhea, symptomatic gonorrhea, pelvic inflammatory disease [PID], disseminated gonococcal infection [DGI], adult conjunctivitis).

Demographic Classification

Data from case reports were abstracted and analyzed for all individuals with gonorrhea who resided in King County at the time of diagnosis. Residence was classified as urban for persons residing within the geographic boundaries of Seattle and as suburban for persons living in incorporated areas of King County or in census tracts with population densities of more than 2000 persons per square mile (based on 1980 US Census statistics). The remainder were classified as rural King County residents. Race and ethnicity were classified according to the CDC case report form 73.688 as White, Black, Hispanic, Asian/Pacific Islander, Native American/Eskimo, or other. Hispanics were classified separately, regardless of race designation.

Socioeconomic Categorization of Census Tract of Residence

Census tracts in Seattle were divided into four socioeconomic categories on the basis of the following population characteristics summarized from the 1980 US Census on Population and Housing[6]: (1) median family income; (2) percent that completed high school; (3) percent employed in professional, managerial, or technical occupations; (4) housing cost index; and (5) percent in owner-occupied housing. A score was computed for each tract from the sum of the tract rankings according to the stated characteristics, and these scores were used to categorize tracts into low, lower-middle, upper-middle, and high SES groupings. These four groupings comprised 25.5%, 28.2%, 18.7%, and 27.6% of the city population, respectively.

Calculation of Gonorrhea Incidence

Incidence rates for gonorrhea were calculated using the total number of cases reported to the Seattle-King County DPH in the 2-year period 1986 to 1987, divided by the total county population count from the 1980 US Census of King County population. This was halved to arrive at an annual rate. Incidence rates by gender, age, race or ethnicity, and census tract of residence were calculated in the same way, using 1980 US Census Bureau data for King County.[6]

Calculation of Estimated Incidence Rates by Marital Status

Marital status was not known for 50% of cases reported from sources other than DPH clinics. In order to estimate incidence rates for marital status, the distribution of marital status for cases where this was known was applied to the unknowns.

Calculation of Adjusted Incidence Rates for Sexually Experienced Women

Gonorrhea rates were calculated for women of White and other races using as denominators national estimates of the number of women who were sexually experienced, as reported in the 1982 National Survey of Family Growth Report;[7] these estimates were adjusted for the number of women in King County in 1980 by age and race.

Statistical Analysis

Correlation between incidence rates and population density in census tracts was examined statistically by computing the correlation coefficient (r).

Results

Incidence of Gonorrhea by Age, Gender, Race, Residence, and Marital Status

Gender and Age Group. The overall annualized incidence rates of gonorrhea by gender were 293 per 100 000 for men and 286 per 100 000 for women. As summarized in Figure 1, the incidence of gonorrhea was highest in females aged 16 to 19 years. For ages 16 to 17 years, incidence rates of gonorrhea were four times higher for females than for males. In contrast, gonorrhea incidence among men was more broadly distributed across age groups, but was higher for men than for women above age 24. Men and women in the 22- to 24-year-old age group had similar rates of gonorrhea.

Race and Ethnicity. Of the 7342 reported gonorrhea cases, 39.6% involved Whites, 48.8% Blacks, 3.3% Native Americans, 3.5% Asians and Pacific Islanders, and 4.9% Hispanics. The annual incidence of gonorrhea per 100 000 by race or ethnicity is shown in Table 1.

There were remarkable differences among racial and ethnic groups in the age and gender distribution of gonorrhea incidence. Among men, the peak incidence of gonorrhea per 100 000 was observed at age 19 for Blacks (11 266) and Native Americans (2299), age 20 for Asians and Pacific Islanders (807), age 21 for Whites (283), and ages 22 to 24 for Hispanics (2597). Among women, the peak incidence per 100 000 was observed at age 18 for Blacks (12 848), Whites (729), and Hispanics (906), and age 19 for Native Americans and Eskimos (6250) and Asians and Pacific Islanders (1068). The overall male:female ratio for gonorrhea incidence was 1.18 for Blacks, 0.73 for Whites, 0.48 for Native Americans, 1.50 for Asians and Pacific Islanders, and 3.51 for Hispanics. As shown in Figure 2, the gender ratio for gonorrhea incidence followed a similar pattern for Whites and Blacks, with females predominating at younger ages and males predominating at ages 25 to 29 years

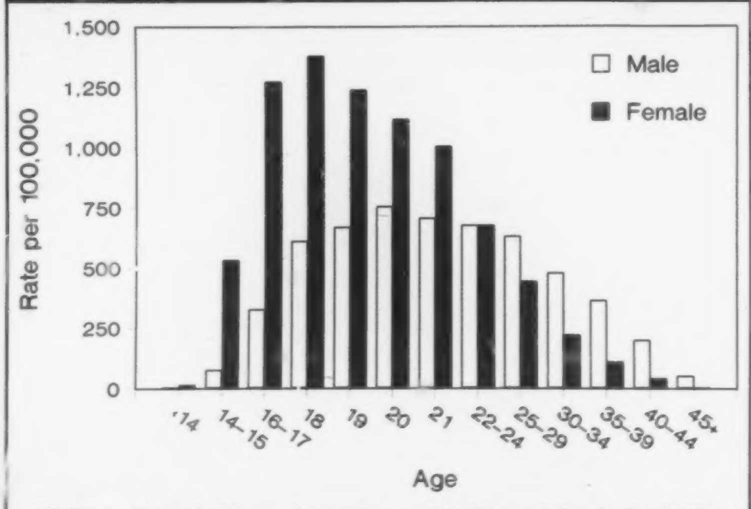

FIGURE 1—Annual incidence of gonorrhea per 100 000 population for Seattle-King County residents by gender and age group, 1986–1987.

TABLE 1—Annual Incidence of Gonorrhea per 100 000 Urban Residents of Seattle According to Race or Ethnicity and Socioeconomic Status (SES), 1986 and 1987

Race or Ethnic Group	SES of Residence Census Tract				All SES Groups
	Low	Lower-Middle	Upper-Middle	Upper	
White	386	153	104	78	121
Asian	310	214	181	125	190
Hispanic	1796	527	427	302	617
NA[a]	1961	1003	640	519	843
Black	3394	3018	2305	1990	3033
All race or ethnic groups	1147	421	224	183	289

[a] NA=Native American.

and above for both groups. However, for Hispanics, rates for females exceeded rates for males only for ages 17 years and under and were about threefold to sixfold lower after age 21. Similarly, for Asians and Pacific Islanders, the incidence in males exceeded that for females for all ages after age 19 and remained low thereafter. In contrast, for Native Americans the incidence rates for females were higher than incidence rates for males for all ages up to 35 years. These differences are large, but are based on relatively small numbers of cases involving Hispanics, Asians and Pacific Islanders, and Native Americans and Eskimos. The differences could also be influenced by underreporting of male cases, as discussed below.

Marital Status. The annual adjusted incidence of gonorrhea per 100 000 was 42 for married men; 67 for divorced, separated, or widowed men; and 1018 for never-married men. The incidence per 100 000 was 33 for married women, 21 for divorced, separated, or widowed women; and 1275 for never-married women. We did not adjust further for marital status in subsequent analyses because data on marital status were not available for all cases.

Urban, Suburban, and Rural Residents. Of the 7342 gonorrhea cases included in this analysis, 74% occurred in urban (i.e., Seattle) residents, 19% in residents of suburban census tracts with ≥2000 residents per square mile, and 7% in residents of rural census tracts (<2,000 residents per square mile). The annualized incidence rates per 100 000 were 555 for urban residents, 143 for suburban residents, and 86 for rural residents.

Socioeconomic Status. In Seattle, which we defined as urban, the population living in low SES census tracts represented only 25.5% of the population but accounted for 58% of the reported gonorrhea. There were 2870 cases in 125 143 residents of low SES census tracts, 1164 cases among 138 294 residents of lower-middle SES census tracts, 409 cases among 91 407 residents of upper-middle SES census tracts, and 494 cases among 135 262 residents of high SES census tracts.

Annual incidence rates according to race or ethnicity and SES are shown in Table 1. Within each racial or ethnic group, there was an inverse and statistically significant relationship between gonorrhea incidence and the SES of the census tract of residents ($P<.01$ for each group). Within each SES grouping, the relative incidence rates by race and ethnicity generally corresponded to the unadjusted relative rates, but the magnitude of the differences by race and ethnicity was reduced. For example, the overall incidence rate was 25 times greater for Blacks than for Whites, but among those of lowest SES the rate was 8.8 times greater for Blacks than for Whites.

Incidence in Female Residents of Low SES Urban Census Tracts. Because the highest incidence of reported gonorrhea occurred in urban census tracts of low SES, and because the completeness of surveillance for laboratory-confirmed gonorrhea may have been highest for females (see Discussion), we further analyzed the incidence rates among women within these census tracts according to age and race or ethnicity. As shown in Figure 3, the highest incidence rates were observed in Black female teenagers, among whom the highest annual incidence reached 17 040 per 100 000 at age 18 and declined thereafter.

Correlation of Gonorrhea Distribution and Incidence with Population Density and Residence. The incidence of gonorrhea by census tract was directly but weakly correlated with population density ($r=.55$, $P=.02$). Mapping of gonorrhea incidence indicated two clusters of high-incidence census tracts, each with ≥1000 cases of gonorrhea per 100 000 population per year, in central and south Seattle (Figure 4). These clusters corresponded to those tracts with the lowest SES rankings and with the highest proportions of non-White residents.

Estimated Incidence Rates in Sexually Experienced Women. By adjusting the denominators to represent only the estimated number of women who were sexually experienced,[7] as advocated by others,[8] the greatest effect on estimated incidence rates was observed, as expected, among teenagers. Regardless of race, adjusted incidence rates increased by about four- to sixfold in 15-year-old adolescent women and by 1.2 to 2.9 in groups of 16- to 17-, 18-, and 19-year-old women (Table 2).

Distribution of Cases by Provider Source. The distribution of gonorrhea by gender, age, and race or ethnicity was compared for cases reported from (1) DPH clinics; (2) hospitals, community clinics, and private physicians; and (3) private family planning clinics. Public clinics reported 56% of cases in males and 38% of cases in females. DPH clinics reported 991 (68%) of 1457 cases in Black men aged over 20 years, vs 545 (58%) of 942 cases in White men of the same age. Otherwise, the distribution of cases by age, sex, and race was not substantially different between DPH clinics and the other groups of providers.

Complications of Gonococcal Infection. Of all women with gonorrhea, 18% received a clinical diagnosis of PID, including 20% of Native Americans, 19% of Whites, 18% of Blacks, and 8% of Hispanics. Fifteen cases of DGI were reported, representing 0.2% of all gonorrhea cases; women accounted for 12 (80%) of these 15 cases. Six cases of adult gonococcal conjunctivitis were diagnosed during the 2-year period. Data were inadequate for estimating the incidence of epididymitis.

Discussion

The distribution of gonorrhea incidence in King County varies markedly according to age, gender, race or ethnicity, and residence. Among individuals classified by age and gender, by far the highest incidence was observed for 18-year-old women, and among adolescents, the ratio of female to male cases was highest for those 14 to 17 years of age. This fact would be obscured by aggregating data into 5-year age groups as in national data sets for sexually transmitted diseases. In explaining these differences in gonorrhea incidence in adolescent males and females, the relative contribution of differences in exposures and differences in host susceptibility remains to be determined. For men and women, the incidence of gonorrhea is highest among certain ethnic minorities living in urban, lower SES census tracts.

A critical feature of this analysis is the adequacy of surveillance. Was case detection representative and complete, or was it incomplete and biased? Evidence for completeness in surveillance includes

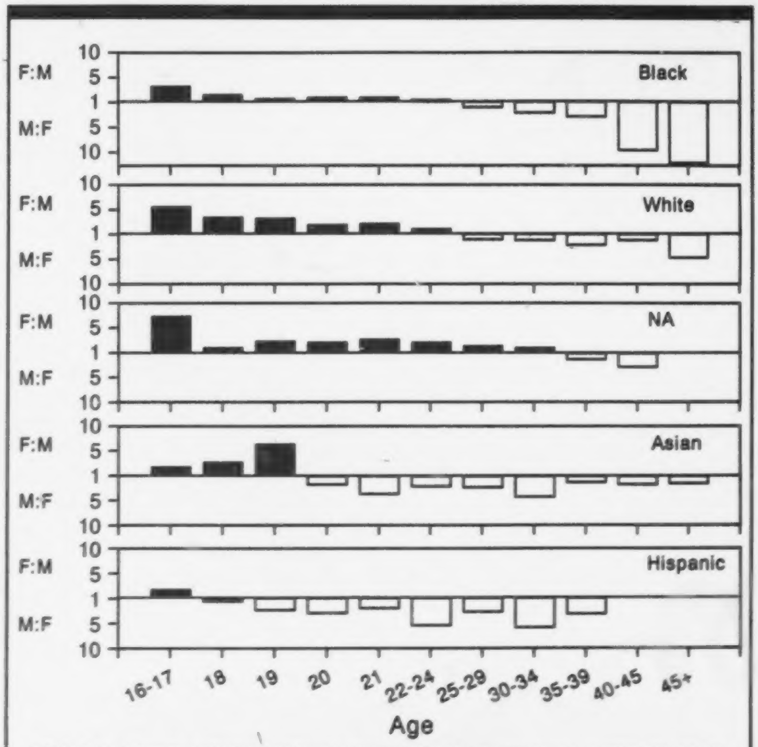

FIGURE 2—Gender ratio for incidence of gonorrhea by age and race or ethnicity. Solid bars above the line indicate a higher incidence in females than in males for that age and race or ethnic group; open bars below the line indicate a higher incidence in males than in females. NA = Native American Indian.

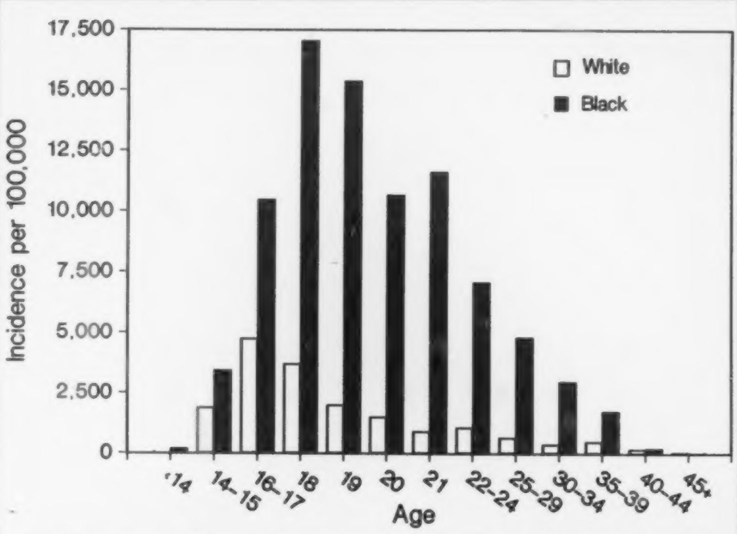

FIGURE 3—Annual incidence of gonorrhea per 100 000 female residents of lower socioeconomic status census tracts in urban Seattle, by age and race (White vs Black), 1986–1987.

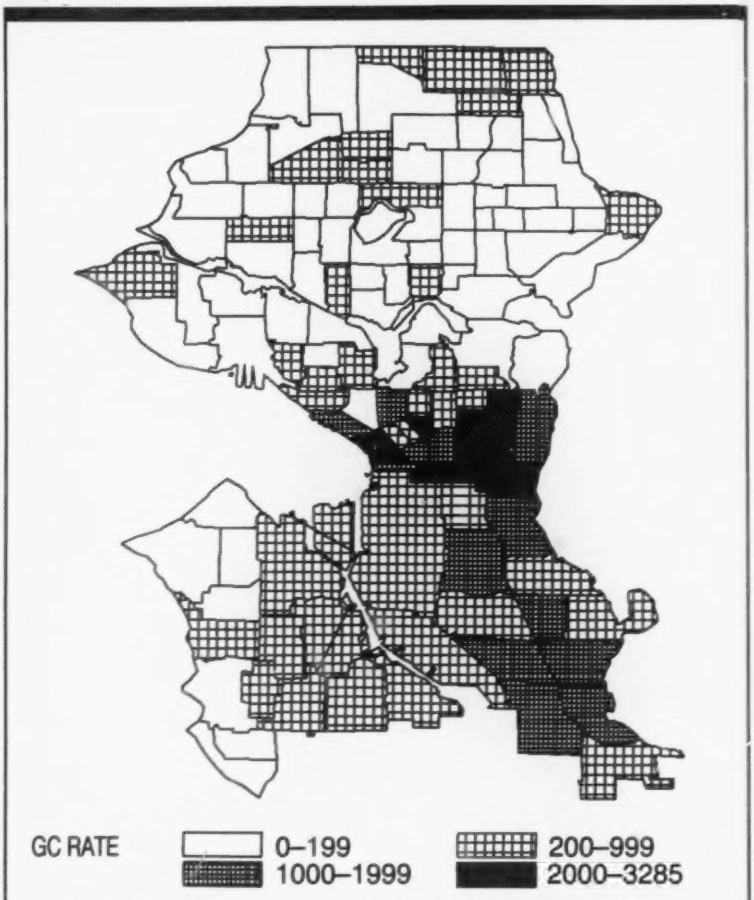

FIGURE 4—Map illustrating the geographic distribution of gonorrhea incidence rates (GC Rate) by census tract of residence in urban Seattle, 1986–1987.

TABLE 2—Gonorrhea Incidence Rates per 100 000 for White and Black Women in King County by Age, Unadjusted and Adjusted for Sexual Experience[a]

	White			Black		
Age	Unadjusted Incidence	% Sexually Experienced	Adjusted Incidence	Unadjusted Incidence	% Sexually Experienced	Adjusted Incidence
14	244	—	—	2395	—	—
15	367	17.9	2051	3873	28.0	13833
16–17	630	35.1	1796	11810	50.0	23621
18	729	54.8	1331	12849	77.0	16686
19	696	69.0	1000	11173	82.0	13625
20	528	84.5	624	10320	93.2	11073
21	590	84.5	698	8687	93.2	9321
22–24	357	84.5	423	6400	93.2	6867

[a]Crude and adjusted gonorrhea incidence rates did not change substantially among women aged ≥ 25 years.[9]

the one-to-one male to female ratio of reported cases; the similar distribution by age, gender, and race or ethnicity of cases from various reporting sources, with the exception noted; and the fact that 53% of cases (44% of males and 62% of females) were reported from non-DPH sources. Nationally, for 1987, only 32% of cases of gonorrhea were reported by private providers.[3] Further, virtually all community clinics, which are partially supported by DPH, use the DPH laboratory for gonorrhea culture diagnosis. Almost all major hospitals and their associated clinics in King County reported cases of gonorrhea through their laboratories during the study period, and these laboratories as well as most commercial laboratories submitted gonococci to the health department laboratory for confirmation and determination of antimicrobial susceptibility.

Regarding reporting of gonorrhea in males, in a unique study of gonococcal urethritis in King County that involved active surveillance of male patients of a sample of private physicians during 1974 and 1975, it was estimated that 38% of cases of gonococcal urethritis were seen in DPH clinics and 62% by all other reporting sources.[9] Some cases in men seen by non-DPH providers were identified by Gram-stained smear of urethral exudate. The proportion of cases of gonorrhea in men that was seen by non-DPH providers in the present study (44%) was lower than the 1974–1975 estimate, and we assume that some cases of gonorrhea in men are still diagnosed by Gram stain, so that no gonococcal isolate would have been sent to the DPH laboratory. It is thus probable that the proportion of cases of unreported gonorrhea seen in the private sector is somewhat higher for men than for women. Thus, we are most confident of the surveillance data for women. It is likely that some of the differences between racial and ethnic groups in the female-to-male ratio for gonorrhea incidence rates could be influenced by underreporting of gonorrhea in men seen in the private sector. However, we think it is unlikely that this could account for the major differences between racial minority groups in the female-to-male ratio and should not account for the steady decline in the female-to-male ratio with increasing age that was observed for most groups.

Even with enhanced surveillance of gonorrhea, cases in males most likely represent incident cases (new infections), while cases in females represent a mixture of incident cases and prevalent cases of uncertain duration: the shorter the duration, the closer the approximation to incident cases. Thus the comparison of male rates to female rates is not straightforward. It is uncertain how age and race or ethnicity influence the duration of infection at the time of diagnosis among women, but again, it seems unlikely that such influences would account for the differences shown in Figure 2.

The 1980 census data for King County were used to estimate the size of the population subgroups at risk during 1986 and 1987. From 1980 to 1986, it is estimated that the total population of King County increased by 7.2%, but that of Seattle increased by only 0.4%. It is possible that unequal trends in the size of the King County population belonging to individual racial and racial or ethnic groups between 1980 and the 1986-1987 period of study could lead to overestimates or underestimates of the calculated incidence of gonorrhea. The Office of Financial Management Forecasting Division, State of Washington,[10] in fact, provides estimates on changes in the size of the population of King County by race and Hispanic origin. These estimates are based on the "component method," which adds births, subtracts deaths, and adds the estimated net migration as estimated from school enrollments.

Because information on Hispanic origin, which is not a category of race, was not collected on vital records or school records, estimates of Hispanic population were derived by taking the 1980 census proportions of persons of Spanish origin in individual racial groups and applying them to the 1986 population estimates for racial groups. Washington State estimates that in King County from 1980 to 1986 the number of Blacks increased by 12%; Whites increased by 5%; Hispanics by 36%; Asians and Pacific Islanders by 26%; and Native Americans by 11%. Thus, the calculated incidence figures by race and Hispanic origin as presented in Table 2 probably represent substantial overestimates for Hispanics and Asians/Pacific Islanders, less substantial overestimates for Blacks, and small overestimates for Native Americans and Whites. Nonetheless, the relative incidence of gonorrhea in these racial and ethnic groups remains essentially as indicated in Table 1.

National data on gonorrhea incidence have not been published by race and ethnicity in recent years. Some states do not report data on race and ethnicity to the CDC for data aggregation and analysis. Nonetheless, in an analysis of gonorrhea incidence in New York State, Rothenberg[11] observed that "core" census tracts accounted for a disproportionately high percentage of gonorrhea cases. These were predominantly inner-city, low-SES census tracts with high density populations of ethnic minorities. It has been theorized that such core group populations are important in maintaining a high endemic level of gonorrhea, and that gonorrhea transmission is strongly affected by behavioral interactions within these population subgroups.[12,13] We observed a very strong inverse correlation of gonorrhea incidence with the SES rank of census tract of residence, despite nonhomogeneity with respect to various measures of SES of individuals within census tracts; this was independent of race and ethnicity. Although large differences in incidence rates between racial or ethnic groups persisted after adjustment for SES, the nonhomogeneity of SES for individuals within census tracts represents a limitation of these analyses.[14] Such nonhomogeneity could help account for the increasing relative difference in the ratio of incidence rates in Whites vs other groups as SES of census tract increased (Table 1).

It is important to attempt to define reasons why the incidence of gonorrhea is so unequally distributed and concentrated within certain core populations. Many factors clearly interact to increase the incidence of sexually transmitted diseases within such core groups.[11-13] Race and ethnicity, in particular, are best viewed not as risk factors or determinants of gonorrhea incidence, but as "risk markers" for behaviors of subgroups within racial or ethnic categories that lead to exposure to or persistence of gonorrhea. Sexual behaviors leading to exposure to gonorrhea include early onset of sexual intercourse,[8] large numbers of casual sex partners, and particularly selection of partners who are themselves members of high-risk core groups.[13] Health behaviors leading to persistence or increased transmission of gonorrhea include failure to recognize symptoms of gonorrhea, delay in seeking treatment after onset of symptoms, continued sexual activity after onset of symptoms, delay in notifying sexual partner(s) of exposure to infection, nonuse of barrier prophylaxis, and noncompliance with therapy. These behaviors in turn may be heavily influenced by high prevalence of or exposure to prostitution, illicit drug use, low economic and academic achievement (including unemployment and school dropout), nuclear family disintegration, limited access to health and behavioral intervention resources, delayed or low utilization of available health care resources, and related socioeconomic and environmental factors. Cultural perceptions of health and disease may also play a role in the uninterrupted cycle of gonorrhea transmission in such settings. Cultural factors have been poorly studied and urgently require more careful behavioral research involving population-based samples.

Overall, these data are consistent with theories that subgroups within the general population have extraordinarily high rates of gonorrhea. Individuals within these groups undoubtedly serve as a pool of high frequency transmitters of infection and play an important role in maintaining the incidence of infection in the general population. The implications of these data are that improved efforts to prevent the transmission of gonorrhea should be directed to high-risk and high-prevalence populations. These efforts may require refocusing and careful evaluation of traditional disease control measures, including culture diagnosis, effective treatment, routine screening of gonococcal isolates for antimicrobial resistance, and partner notification. Appropriate reallocation of resources for personnel and clinics in geographic areas more accessible to high-risk communities may be needed. In addition, behavioral interventions must be developed for high-incidence groups. Primary interventions clearly should be initiated no later than age 13 in order to reach teenagers who may be exposed at a very early age not only to gonorrhea, but also to other sexually transmitted diseases. Secondary behavioral interventions for high-risk individuals who present with such diseases are also needed. Development and careful evaluation of behavioral interventions based on current theoretical models of human behavior deserve the highest priority for research.

In summary, data on gonorrhea transmission and incidence within population subgroups can serve as a useful proxy for high-risk sexual behaviors that have implications for the transmission of other sexually transmitted diseases, including human immunodeficiency virus (HIV) infection. The patterns of incidence of gonorrhea observed in this study are similar to patterns of heterosexual transmission of syphilis, chancroid, and HIV, which, like gonorrhea, appear to be concentrated in urban low-SES population subgroups in the United States.[1,15] These data reinforce the need for a reassessment and refocusing of traditional disease control strategies and for development of age-specific and culturally sensitive behavioral and social intervention strategies. ☐

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Most Physicians Receive Income Guarantees, Survey Confirms

The majority of physicians who accept new positions through hospital recruitment programs receive income guarantees, new survey data show. Among the 151 hospitals surveyed by Atlanta-based Jackson and Coker, 91% offer such guarantees, which include practice start-up or development costs and a minimum income level for the physician in the first year of practice.

"Income guarantees are nothing less than security blankets," said Randy Gott, Jackson and Coker's director of research. In most cases, the physician exceeds the guarantee in the first year of practice." In fact, over half of the respondents reported subsidy payments to recruited physicians of less than 25% of the guaranteed amounts.

So intense is competition for physicians among hospitals that 73% of hospitals surveyed reported that they sponsor or subsidize physician searches for existing members of their medical staff. Among the most popular forms of shared participation are reimbursement of candidates' interview expenses, fees paid to physician recruitment firms, reimbursement of candidates' relocation costs, and income guarantees.

But how hospitals choose to use these incentives depends on several factors, including the hospital's location and the physician's specialty, according to Gott. "While a hospital in a major city might entice a physician with free office rent, reimbursement of relocation expenses, and a small financial guarantee, a rural hospital might need to offer that physician a larger and more extended financial guarantee," he said. In addition, income guarantees offered will vary by specialty and will depend upon the physician's earning potential in a specific region.

Income guarantees vary in duration. Sixty-nine percent of hospitals reported that their guarantees last 1 year; and another 21% said that their guarantees last 2 years. Only 6% of hospitals surveyed offered guarantees of 6 months. Gott is convinced that hospitals will increasingly offer 2-year guarantees to physicians in highly competitive specialties where practices tend to develop more slowly.

While 74% of hospitals reported a payback provision as a part of their income guarantees, 30% of those surveyed said they would waive a portion of the amount to be paid back if the physician stayed in the practice for a designated period of time. For example, if a physician remained in the practice for 1 year beyond the guarantee period, the hospital would forgive 33% of the guarantee. But if the physician stayed for 3 years, the hospital would forgive the entire amount to be paid back.

In 33% of hospitals surveyed, physicians can pay back income guarantees by performing a variety of services, including administrative duties (67%), consulting services (30%), emergency department coverage (27%), teaching continuing medical education classes (27%), and caring for indigent patients (27%). Added Gott: "Nearly all hospitals that require a payback allow a minimum of 1 year. They want to make a payback as easy as possible."

For more information on this survey, contact Randy Gott or Becky Fitzgerald at 800/548-5393.

Public Health Intervention in a Cocaine-Related Syphilis Outbreak

ABSTRACT

Background. Cocaine users and prostitutes are at high risk for syphilis, but disease control is difficult among these populations. During a cocaine-related syphilis outbreak in Chester, Pennsylvania, in 1989, we conducted a control program at sites where sex and drugs were sold.

Methods. During a 2-week period, investigators recruited persons from these sites for interview, serologic testing, and empiric treatment.

Results. Among 136 persons screened, 25 (18%) had early syphilis and 26 others (19%) had recent sexual contact with early syphilis patients. All were treated at initial screening at a cost of $402 and 12 investigator hours per case, compared to $470 and 20 hours per case when treated during routine investigator activities. This program may have contributed to a short-term decline in syphilis incidence in Chester by reducing the period of infectivity of these patients.

Conclusions. Screening and empiric treatment of persons at sites where sex and drugs are sold can be useful in short-term control of cocaine-related syphilis outbreaks. (*Am J Public Health*. 1991;81:1259–1262)

Jonathan R. Hibbs, MD, and Robert A. Gunn, MD, MPH

Introduction

Syphilis case-finding increased in the United States during the 1980s.[1] Transmission among cocaine users and prostitutes contributes to this problem,[2–4] and the disease is difficult to control in these populations. Syphilis control depends on treatment of patients in the early stages.[5–7] The usual approach is to treat all persons with clinical or serologic evidence of infection and their sexual partners.[8]

The Pennsylvania Health Department noted increased early syphilis case-finding in southeastern Pennsylvania during 1988; one town, Chester, had a fivefold increase during 1988 compared with 1987 (Pennsylvania Department of Health, unpublished data). Because epidemiologic information identified cocaine use and prostitution as risk factors, the Health Department targeted control efforts toward locations thought to be used for the sale of sex and drugs. In contrast to the usual approach, empiric treatment was given to all persons at the time of screening.

Methods

Chester is located 20 miles south of Philadelphia. In 1980, 24% of Chester residents lived at or below the poverty level, and 46% of residents had not graduated from high school.[9] Chester's population was 53% African-American in 1980. Public facilities for sexually transmitted disease (STD) treatment in Chester included a hospital emergency room and two publicly supported STD clinics. In 1988, 9% of syphilis cases in Chester were reported from public clinics.

Primary, secondary, and early latent syphilis (referred to collectively as early syphilis) were defined using generally accepted criteria.[10] Involvement with prostitution was defined as giving or receiving money or drugs for sex within the year prior to interview. Drug use was defined as use of any illicit drug within the year prior to interview. A syphilis contact was defined as any person named as the sexual partner of an untreated early syphilis case-patient within 6 weeks before the case-patient interview.

Health providers and laboratories are required to report to the Health Department all cases of syphilis and all laboratory test results indicative of syphilis. A Health Department investigator interviews the patient. If early syphilis is confirmed, the investigator attempts to interview the patient's recent sexual partners and facilitates medical examination and treatment for them and the case-patient. Investigators are responsible for follow-up of all reportable STDs; however, during the time of the screening program, investigator activities in Chester were devoted exclusively to syphilis.

The syphilis screening and treatment program took place between 4 and 11 PM on 6 days between April 17 and 28, 1989. Persons were recruited for screening at community sites in Chester, then taken to a treatment center. Recruiting sites were

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Jonathan Hibbs, at the time of this study, was with the Epidemiology Program Office of the Centers for Disease Control, Atlanta, Ga, and the Bureau of Epidemiology of the Pennsylvania Health Department. He currently is with the National Institutes of Health in Bethesda, Md. Robert A. Gunn is with the Epidemiology Program Office of the Centers for Disease Control, Atlanta, Ga.

Requests for reprints should be sent to Jonathan Hibbs, MD, NIH/NHLBI/CHB, 10/7C114, National Institutes of Health, Bethesda, MD 20892.

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TABLE 1—Cocaine Use, Prostitution, and Early Syphilis among Syphilis Screening Program Clients, Syphilis Intervention Program, Chester, Pa, 1989

Risk Factor	Screening Program Clients	Cases (%)	Rate Ratio (95% CI*)
Cocaine use and prostitution	15	7 (47)	4.3 (1.6–9.4)
Cocaine only	23	6 (26)	2.4 (0.8–6.3)
Prostitution only	6	2 (33)	3.1 (0.4–9.1)
Neither	92	10 (11)	1.0 (reference)
Total	136	25 (18)	

*CI=confidence interval.

TABLE 2—Cost of Syphilis Control Activities in Chester, Pa, April 15–28, 1989, Syphilis Intervention Program

	Syphilis Control Activities	
Factors	Screening and Treatment Program	Routine Investigator Activities
Investigator hours	300	415
Investigator cost	$ 6,914	$9,564
Other costs	$ 3,143	$ 300
Total costs	$10,057	$9,864
Cases treated	25	21
Hours per case treated	12	20
Cost per case treated	$ 402	$ 469

chosen based on information provided by syphilis patients that drugs or sex were sold nearby. Police agreed that they would be absent from these sites during screening. Disease investigators alerted persons frequenting these sites about the screening program several days before it took place. On 3 days, the treatment facility was the Chester State Health Center, and on 3 days it was a mobile van equipped for patient examination and moved within a block of recruiting sites.

Investigators introduced themselves to all persons found at recruiting sites and invited them to be tested and treated for syphilis. If the person agreed, the investigator transported him or her to and from the treatment site. At the treatment site, clients were interviewed, informed of potential side effects, and asked to sign a consent form allowing syphilis testing and treatment. Consenting clients had blood drawn by a nurse, were examined by a physician, and unless they denied any sexual contacts within the past year were treated for early syphilis.[11] Clients receiving penicillin were asked to remain at the treatment site 10 minutes for observation. Rapid plasma reagin (RPR) tests were performed on all blood samples. Reactive samples were analyzed with the fluorescent treponemal antibody test.

Investigator costs were calculated by adding wages, expenses, and benefits of all investigators in Chester during the screening program. In order to identify the fraction of investigator costs due to the screening program itself, the 300 investigator hours spent on the screening program were divided by the 715 investigator hours spent doing syphilis control in Chester during the weeks the screening program took place. Other screening costs included physicians' and nurses' wages, medications, phlebotomy and injection equipment, and van rental.

Routinely identified case-patients treated at public STD clinics cost the Health Department $20 each for medications and clinical fees. Routinely identified case-patients treated by private physicians cost the Health Department only investigator time; the cost of their treatment is unknown. We calculated exact 95% confidence intervals (CI) for rate ratios (RR)[12] and odds ratios (OR).[13]

Results

During the program, 136 clients were screened, of whom 80 (59%) were male and 127 (93%) were African-American. The median age of screening program clients was 20 years (range=11–52 years). The risk of syphilis was greater among female clients (15/56 [27%] vs 10/80 [13%] among male clients; rate ratio [RR]=2.3; CI=1.1–4.7) and among clients 20 years of age or older (21/71 [30%] vs 4/65 [6%] among younger clients; CI=1.7–13.3).

Among 136 persons screened, 38 (28%) reported cocaine use, and 21 (15%) reported involvement with prostitution. Clients reporting these risk factors had increased risk of syphilis (Table 1). Among 34 clients known to be case contacts, 8 (24%) had early syphilis, compared with 17 (17%) of the 102 clients not identified as case contacts (RR=1.4; CI=0.6–3.1). We treated 114 (84%) of the 136 screening program clients, including all case-patients and case contacts. No adverse drug reactions were noted.

The screening program cost the Pennsylvania Health Department $10 057, of which $6914 (69%) was investigator related (Table 2). Compared with routine investigator activities in Chester during the same period, the screening program cost fewer dollars and fewer investigator hours per case treated.

In the first 32 weeks of 1989 (January 1 through August 12), 302 early syphilis cases were diagnosed among Chester residents—1130 cases per 100 000 persons annually. Case-finding peaked in Chester during the weeks of the screening program (Figure).

Personal characteristics of case-patients diagnosed during the screening program were generally similar to those of routinely reported cases, although the screening program case-patients were slightly more likely to report cocaine use (Table 3). Early latent syphilis was more common among screening program case-patients than among routinely reported case-patients (64% vs 44%; OR=2.2, 95% CI=0.9–5.7), particularly among men (90% vs 50%; OR=9.0, 95% CI=1.2–401).

Discussion

We intervened during the peak of an urban syphilis outbreak. Routine partner notification had limited efficacy in controlling this outbreak in its early stages. Disease control was complicated by the involvement of cocaine users and prostitutes in an economically depressed area typical of urban STD transmission.[14] Multiple anonymous sexual contacts and the illicit nature of prostitution and drug use may limit the effectiveness of partner notification among such populations.[15,16] Increasing the number of investigators,[17] broadening criteria for screening,[18] and

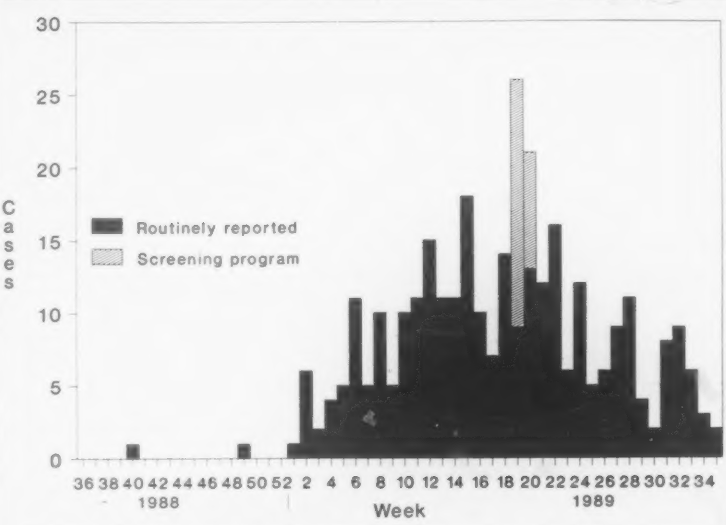

FIGURE 1—Syphilis cases by date of diagnosis and report source, Syphilis Intervention Program, Chester, Pa, 1989.

TABLE 3—Screening Program and Routinely Reported Syphilis Cases, by Age, Gender, Drug Use, and Involvement with Prostitution, Syphilis Intervention Program, Chester, Pa, 1989

	Screening Program Cases (n = 25)		Routinely Reported Cases (n = 277)[a]			
Characteristic	No.	(%)	No.	(%)	OR	(95% CI)[b]
Age <20 years	4	(16)	43	(16)	1.0	(0.3–3.3)
Female	15	(60)	157	(57)	1.2	(0.5–2.9)
Cocaine user[a]	13	(52)	82	(35)	1.7	(0.7–4.2)
Use of any illicit drug by case-patient or sexual partners[a]	16	(64)	122	(52)	1.3	(0.5–3.3)
Prostitutes and clients[a]	9	(36)	61	(26)	1.4	(0.5–3.5)

[a] Drug and prostitution information available for only 234 routinely reported cases.
[b] OR=odds ratio; CI=confidence interval.

targeting areas where high-risk persons congregate[19,20] may be indicated in STD outbreaks that routine partner notification fails to control. Our intervention combined these elements and added on-site empiric treatment.

We showed that such a program can treat infected persons and their partners in high-risk areas. In a similar screening program done during evening hours in Philadelphia, three residential sites and a soup kitchen were used. Among the 268 persons screened, 12 (4.5%) had untreated early syphilis (M. Goldberg, written communication, 1990). In a daytime outreach program conducted at recreation centers, development agencies, and a food give-away center in Rochester, NY, 205 persons were screened and only 1 (0.5%) had untreated early syphilis (T. Huntley, written communication, 1990). A high yield of cases may depend on identifying appropriate sites and times for the screening program. We believe the training of disease investigators, their experience in Chester, and their ability to establish rapport with persons in high-risk areas also contributed to the success of our program.

Our screening program was less costly per case treated than routine investigator activities in Chester during the same period. Our estimate of routine activity costs is probably low, because we excluded the cost of private treatment for routinely identified case-patients.

This program's impact on the outbreak is difficult to substantiate. The number of cases detected per week declined after our program, suggesting an effect on incidence. It is possible that cases identified during the screening program were identified earlier than they would have been otherwise, and all the program accomplished was earlier case-finding. If that is true, the program decreased the period of infectivity of case-patients and case contacts, which should reduce STD transmission.[21] The high percentage of early latent cases (which are asymptomatic) among our male case-patients also suggests that the screening program reduced transmission, because asymptomatic men account for a disproportionate share of STD transmission[21] and are unlikely to seek medical attention on their own.

Empiric on-site treatment was a new facet of our program. Such treatment "pits the clinical importance of accurate diagnosis against the policy of public health prevention." (p. 1081)[8] Fortunately, recommended treatment of early syphilis in most patients[11] is cheap, convenient, and has few major side effects. We found empiric treatment to be cost beneficial and well tolerated. We believe that providing such treatment for high-risk populations is safe, reaches persons who may not return for follow-up, and may reduce syphilis transmission.

The mobile van in this program proved to be an effective addition to a standing permanent structure. The usefulness of this strategy is influenced by proximity to a permanent clinic and availability of an adequate, reasonably priced mobile unit. Mobile units could also improve long-term contact between health workers and the community; however, our program did not evaluate the use of mobile units for this purpose.

Since treatment does not change underlying STD risk factors,(3) interventions of this kind cannot provide long-term disease control. For short-term reductions of syphilis incidence in outbreaks associated with cocaine and prostitution, however, an approach such as ours is recommended, especially the use of empiric first-visit treatment of persons recruited for screening. Consideration should be given to using such screening for longer periods, repeating the program in the same population at intervals to determine whether a change in incidence occurs, and adding other sites. If the case yield is high, screening programs should include empiric treatment or on-site RPR screening followed by immediate therapy. Only by using a variety of strategies can the best ones be selected to combat urban syphilis outbreaks. Prompt development of such strategies is important, particularly in light of the potential for human immunodeficiency virus spread among persons with ulcerous STDs.[22] ☐

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Diverging Gonorrhea and Syphilis Trends in the 1980s: Are They Real?

Kenneth A. Gershman, MD, MPH, and Robert T. Rolfs, MD

ABSTRACT

Background. The purpose of this study was to evaluate whether the divergence in national trends of gonorrhea and syphilis from 1986 to 1989 in the United States was real and if overall trends masked a contemporaneous increase in both diseases in a core group.

Methods. We analyzed the following: (1) reported cases of gonorrhea and primary and secondary syphilis in the United States for the years 1981 to 1989, (2) gonorrhea screening results from six states for the years 1985 to 1989, and (3) reported cases of gonorrhea and primary and secondary syphilis by census tract for the years 1986 to 1989 in one city.

Results. The incidence of gonorrhea decreased 22% in the United States from 1986 to 1989 while the incidence of primary and secondary syphilis increased 59%. Among Blacks, syphilis incidence increased 100% and gonorrhea incidence decreased 13%; among Whites and Hispanics, the incidence of both diseases decreased. Results from gonorrhea screening among females in six states agree with gonorrhea incidence trends in those areas. Race-specific and census tract analyses of data from a number of metropolitan areas where overall rates diverged did not demonstrate a group in which the incidence of both diseases increased.

Conclusions. We conclude that diverging trends of gonorrhea and syphilis from 1986 to 1989 are real and emphasize differences in the epidemiologic characteristics of these two sexually transmitted diseases. (*Am J Public Health*. 1991;81:1263–1267)

Introduction

Declines in numbers of reported cases of gonorrhea and syphilis in the first half of the 1980s suggested that Americans had changed their sexual behavior in response to the AIDS epidemic.[1–3] Since 1986, however, syphilis incidence has increased[4] while gonorrhea incidence has continued to decline.

The recent spread of syphilis has been associated with drug use, especially cocaine.[5–8] Drug use has also been associated with the spread of gonorrhea, but this association has been less extensively studied.[9–11] Since syphilis and gonorrhea are transmitted in the same way, an increase in gonorrhea incidence, rather than a decrease, might be expected during the time when syphilis incidence increased.

The decrease in gonorrhea incidence, although probably real, could be the result of decreased surveillance and case detection activities. Anecdotally, reductions have occurred in sexually transmitted disease (STD) clinic visits in some cities and in gonorrhea screening programs in many states. Such decreases in gonorrhea case detection activities could result in artifactual declines in gonorrhea incidence. Alternatively, an overall gonorrhea decrease might mask an increase in the same core group in which syphilis has increased. To better understand if the divergence in national trends of gonorrhea and syphilis is real and whether this divergence masks a contemporaneous increase in both diseases in a subgroup of the population, we analyzed several types of data. These included (1) reported cases of gonorrhea and primary and secondary (PS) syphilis for the years 1981 to 1989, (2) gonorrhea screening results from six states for the years 1985 to 1989, and (3) reported cases of gonorrhea and PS syphilis by census tract for the years 1986 to 1989 in a city experiencing a large increase in syphilis.

Methods

We used summary data from 1981 to 1989 on cases of gonorrhea and PS syphilis reported to state health departments and sent annually to the Centers for Disease Control (CDC). Annual reports included summary data by gender, racial or ethnic group, age group, and source of report (public, private, or military) for each state and six large cities (Baltimore, Chicago, Los Angeles, New York City, Philadelphia, and San Francisco). Cases reported from public and private sources were combined; military cases were not included in these analyses. Racial or ethnic group was categorized as Black (not of Hispanic origin), White (not of Hispanic origin), Hispanic, Asian or Pacific Islander, or American Indian or Alaska Native.

Incidence rates for the United States were calculated using US Bureau of the Census population estimates for 1981–1989.[12,13] Estimates of the White and Black Hispanic populations were subtracted from estimates of Whites and Blacks to calculate estimates for Whites not of Hispanic origin and Blacks not of Hispanic origin.

Five states and two cities in 1983 and two states and three cities in 1984 did not

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The authors are with the Division of STD/HIV Prevention, Center for Prevention Services, Centers for Disease Control.

Requests for reprints should be sent to Information Services, Center for Publication Services (EO6), Centers for Disease Control (EO2), Atlanta, GA 30333.

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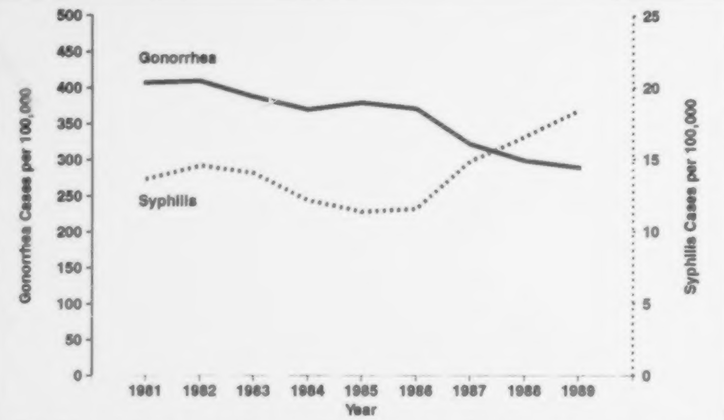

FIGURE 1—Trends in incidence of gonorrhea and primary and secondary syphilis, United States, 1981–1989.

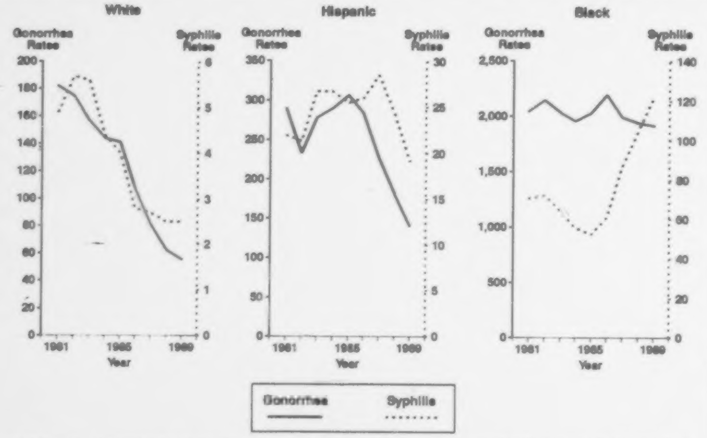

FIGURE 2—Trends in race-specific incidence of gonorrhea and primary and secondary syphilis, United States, 1981–1989. Rates = No. of cases per 100 000 persons.

submit annual reports of gonorrhea cases to CDC. Estimates for these areas were projected by averaging the sex and race distributions of cases in each area from the years before and after the missing year(s). Similar methods for interpolating missing syphilis data have been described previously.[4]

The proportion of gonorrhea cases reported with race not specified varied from year to year (17–19% for the years 1981 and 1986–1989 and 22–26% for 1982 through 1985). No cases were reported with sex not specified. Cases with race not specified were prorated to the race and sex distribution of known cases. To assess the validity of prorating, we tested varying assumptions, including the extreme assumptions that unspecified cases were all Black or all White. Gonorrhea trends were not sensitive to these extremes. Similar methods for prorating cases of syphilis with race not specified have been previously described.[4]

For the analysis of state trends, an increase in syphilis incidence was defined as a greater than 33% rate increase from 1986 to 1989 with both 1988 and 1989 rates higher than the 1986 rate. Only states with at lease 25 PS syphilis cases in 1989 were included.

Results of gonorrhea screening in women for the years 1985 to 1989 were obtained from six state STD programs where these data were available. The numbers of gonorrhea cultures performed and the numbers positive by type of facility were reported. STD clinic results were excluded and overall annual positivity rates (percent positive) were calculated for all other sources combined.

Cases of gonorrhea and PS syphilis by census tract for Rochester, NY, for the years 1986 to 1989 were obtained from the New York State STD program. Rochester was selected because gonorrhea and syphilis trends diverged there and census tract data were available. Census tracts were ranked by the number of PS syphilis cases reported in 1989. Core syphilis census tracts were defined as those with the highest ranking that cumulatively accounted for approximately 50% of cases.[14] Incidence rates were calculated using 1980 US Bureau of the Census population estimates by census tract.[15]

Results

National Trends

The incidence of both gonorrhea and syphilis decreased in the United States from 1981 to 1985 (Figure 1). From 1986 to 1989, the incidence of gonorrhea decreased 22% from a rate of 370.8 to 289.1 cases per 100 000 persons. The incidence of PS syphilis, however, increased 59% from a rate of 11.6 to 18.4 cases per 100 000 persons.

Among Whites, the incidence of gonorrhea decreased 48% from 1986 to 1989 (107.1 to 56.4 cases per 100 000), while that of syphilis decreased 11% (2.8 to 2.5 cases per 100 000) (Figure 2). Among Hispanics, the incidence of gonorrhea decreased 50% during this time (283.5 to 140.9 cases per 100 000) and syphilis incidence decreased 27% (26.0 to 19.1 cases per 100 000). However, among Blacks, a marked divergence in trends occurred: gonorrhea incidence decreased to a lesser extent than among the other groups, 13% (2197.1 to 1912.4 cases per 100 000), while syphilis incidence increased 100% (61.0 to 121.8 cases per 100 000). Sex-specific trends within each racial or ethnic group were generally similar to overall racial or ethnic trends. While syphilis incidence increased among White females from 1986 to 1989 (from 1.4 to 1.9 cases per 100 000 persons), their rates remained extremely low. Syphilis incidence among Hispanic females did not change appreciably during this time period.

Only 18 states reflect the divergence in national trends of gonorrhea and syphilis from 1986 to 1989. Although gonorrhea incidence rates decreased or remained unchanged in 45 states, syphilis incidence rates increased in only 22 states. In 18 of these 22 states, gonorrhea incidence decreased or remained unchanged.

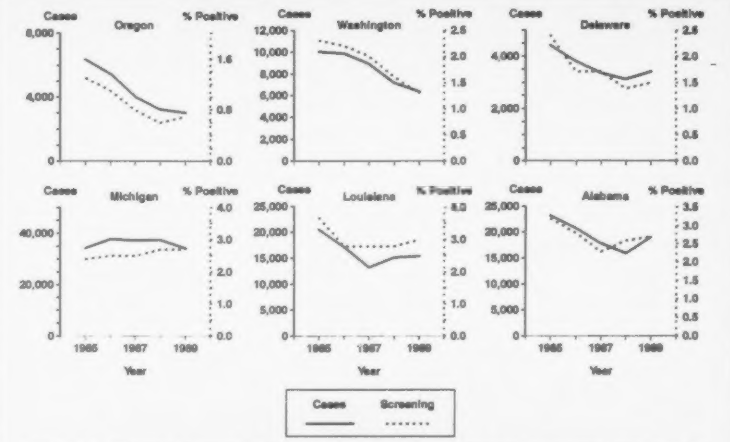

FIGURE 3—Trends in reported cases of gonorrhea and gonorrhea screening positivity rates (% positive cultures) in women, selected states, 1985–1989.

TABLE 1—Incidence of Gonorrhea and Primary and Secondary Syphilis in Core and Noncore Census Tracts in Rochester, NY, 1986–1989[a]

	Core		Noncore	
Year	Gonorrhea	Syphilis	Gonorrhea	Syphilis
1986	3721	11	1001	6
1987	3202	22	835	6
1988	4079	175	1049	35
1989	3867	266	1099	38

[a] Incidence is cases per 100 000 persons; core census tracts are those with 56% of primary and secondary syphilis cases in 1989.

Gonorrhea Screening

To evaluate whether changes in surveillance methods or case detection may account for the decrease in gonorrhea incidence, we compared trends in screening positivity rates in women and reported cases of gonorrhea in six states in which syphilis has increased (Figure 3). Screening and reported case trends were quite similar between 1985 and 1989 in these states. In two states (Oregon, Washington), screening positivity rates and reported cases both declined throughout the period 1985 to 1989; in three states (Alabama, Delaware, Louisiana), both trends initially declined then increased in parallel; and in one state (Michigan), both trends remained relatively unchanged. For two of the six states, the numbers of women screened decreased by 5% or less from 1986 to 1989. In the other four states, the numbers of women screened declined by 15% to 51%.

Core hypothesis. To evaluate whether the overall divergence in trends of gonorrhea and syphilis masked a contemporaneous increase in both diseases in a core group, we first analyzed race-specific case totals for 1986–1989 from several metropolitan areas where these data were available. In four metropolitan areas (Seattle, Memphis, Charlotte, Los Angeles) with an overall divergence in trends, stratification by race did not identify a group in which the incidence of both gonorrhea and syphilis increased.

To further evaluate the core concept, we analyzed census tract data from one metropolitan area. In Rochester, NY, gonorrhea incidence remained relatively constant from 1986 to 1989 (12% increase), while syphilis incidence increased 11-fold. We compared gonorrhea and syphilis incidence rates in the census tracts with approximately 50% of the city's reported syphilis cases (the core area) in 1989 (Table 1). Core incidence rates for syphilis and gonorrhea were 7.0 and 3.5 times higher, respectively, than noncore rates in 1989. Even in the core syphilis census tracts, however, trends for the two diseases diverged as syphilis incidence rates increased 24-fold, while gonorrhea incidence rates showed little change. A similar divergence, but of lesser magnitude, also occurred in the noncore census tracts. Of note, gonorrhea incidence rates for Rochester in 1987 were substantially lower than rates for all other years from 1984 to 1989.

Discussion

From 1986 to 1989, national trends of gonorrhea and syphilis incidence diverged: gonorrhea declined while syphilis increased markedly. This divergence occurred among Blacks, but not among Whites or Hispanics. Results from gonorrhea screening among females in a number of states support the overall decrease in gonorrhea incidence as real, rather than the result of changes in surveillance. We were unable to demonstrate a core group in which the incidence of both gonorrhea and syphilis increased.

Gonorrhea screening positivity rates declined from 1985 to 1989 in five of the six states we assessed, as did gonorrhea incidence. In the sixth state, both gonorrhea screening and incidence trends changed little. Screening positivity rates should be a reliable indicator of gonorrhea prevalence if the population screened remains relatively constant over time. For the two states in which the numbers of women screened changed little, parallel decreases in both screening and incidence trends strongly suggest that a real decrease in gonorrhea incidence occurred. In the other four states, however, the numbers of women screened decreased substantially from 1985 to 1989. Declines in screening positivity rates in these states may reflect real decreases in gonorrhea prevalence or may be the result of screening fewer high-risk women. Since STD programs are likely to reduce screening at clinics with the lowest yield and maintain screening among high-risk groups, these trends probably reflect true decreases in gonorrhea incidence.

Geographic clustering of both gonorrhea and syphilis cases among "core" census tracts in urban areas has been demonstrated previously.[14,16,17] Differential

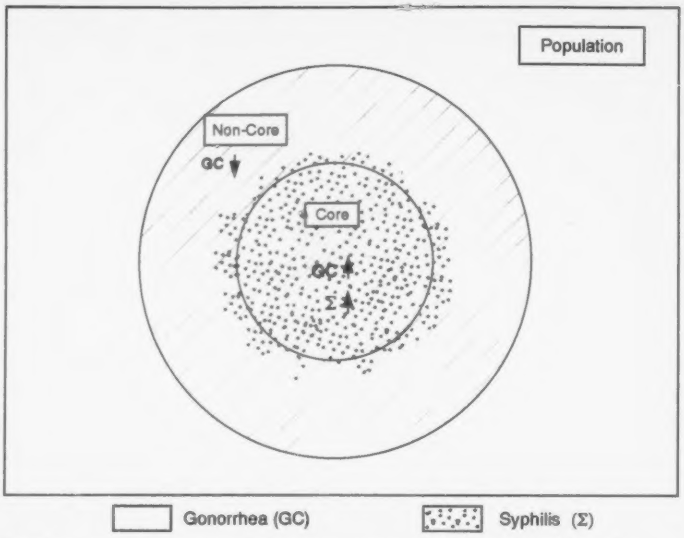

FIGURE 4—Core model of distribution of gonorrhea and syphilis in a metropolitan area.

changes in disease transmission for core and noncore groups might mask trends in the smaller core group (Figure 4). For example, the exchange of sex for money or drugs in a core group might lead to increases in both gonorrhea and syphilis incidence. Screening plus condom use or other behavior change, however, might decrease gonorrhea incidence in the much larger noncore group and mask the core group increase. Because the distribution of syphilis is largely confined to a core group, overall syphilis trends would be those of the core group.

If transmission of both infections increased in the core group while decreasing in the general population (noncore group), an overall decrease in gonorrhea incidence and increase in syphilis incidence would occur.

Neither the analysis of race-specific data in several urban areas with an overall divergence, nor the census tract analysis of another city demonstrated a group in which the incidence of both diseases increased. Race and geography are almost certainly not direct risk factors for STD; presumably they are surrogate markers for behavioral factors, which if better defined and measured would allow us to more clearly identify the core group. Race and geography alone may be inadequate to define a core group in which the incidence of both gonorrhea and syphilis increased.

It seems likely, however, that transmission of *Neisseria gonorrhoeae* has not been affected to the same extent by whatever factors have increased the spread of syphilis. Several factors might account for differences in the epidemiologic patterns of gonorrhea and syphilis in the United States. First, gonorrhea and syphilis may have different core groups. Although studies suggest that the core group involved in the recent syphilis increase is the same as the crack-using population, less evidence exists that these two groups overlap with the gonorrhea core group.[18] Second, gonorrhea and syphilis may have the same core group, but the factors that have increased the spread of syphilis may not affect the spread of gonorrhea. Oral sex has been reported to be the most frequent type of sexual activity involved in the exchange of sex for crack[19] (S. K. Koester, Ph.D., November 1990, written communication), and oral sex might be more effective in spreading syphilis than gonorrhea. Transmission of pharyngeal gonococcal infection to sex partners is thought to be inefficient and relatively rare[20] and, therefore, would be unlikely to sustain gonorrhea spread. Third, differences in health care seeking behavior among those with gonorrhea and syphilis might also contribute to the divergence. Males with gonococcal urethritis tend to seek treatment shortly after the onset of symptoms.[21] Prompt treatment translates into a short infectious period and interrupted transmission. In contrast, a large proportion of persons with infectious syphilis do not seek medical attention despite signs and symptoms.[22] Crack use and related factors such as unemployment and poverty may contribute to delays in seeking health care among persons with syphilis.

From 1986 to 1989, gonorrhea and syphilis trends have diverged for Blacks, among whom syphilis has increased dramatically since 1986, but not for Whites or Hispanics. An understanding of why syphilis but not gonorrhea incidence has increased in this population will better define transmission dynamics of these two diseases and, therefore, contribute to more successful control strategies. ☐

Acknowledgments

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Hopkins Announces New Vaccine Effective Against Life-Threatening Infection

Johns Hopkins School of Public Health researchers have proven a new Hib (Haemophilus influenzae type B) vaccine effective in preventing life-threatening infection in young children when administered at 2 and 4 months of age. The vaccine, called Hib-OMPC, was tested on Navajo children who, like some other Native American populations, are 10 to 50 times more susceptible to Hib-related disease than the general US population.

"The widespread use of Hib-OMPC vaccine could virtually eliminate *H influenzae* and its deadly effects in children around the world," said lead author Mathuram Santosham, MD, MPH, director of the Johns Hopkins Center for American Indian and Alaskan Native Health. The results of the trial are reported in the June 20 *New England Journal of Medicine*.

Infection with Hib—which is actually a bacterium, not a "flu" virus—strikes 20 000 children in the United States per year. In half the cases the infection develops into meningitis, a potentially fatal infection of tissues surrounding the brain and spinal cord. Five to 10 percent of the infected die; survivors often develop neurological disorders, including hearing loss, paralysis, and speech problems. In addition to being the leading cause of meningitis, Hib can cause pneumonia; blood poisoning; infections of the ears, joints, skin, and lining of the heart; and a potentially fatal throat infection.

"The incidence of Hib-related infection is extremely high in Native American children because this population has a lower antibody response to Hib vaccines than the general population. We learned these differences in a recent study on Apache children," Santosham said.

"Proving the efficacy of Hib-OMPC vaccine in this high-risk population should also predict its efficacy in other populations," he added.

With the help of Raymond Reid, MD, MPH, a Hopkins research associate who is a Navajo physician, and a number of other investigators, Santosham directed the double blind placebo-controlled clinical trial in 3486 Navajo infants from July 1988 to August 1990 in Whiteriver, Az. The Hib vaccine was shown to be effective in preventing *H influenzae* disease when two doses were administered at 2-month intervals beginning at 6 to 12 weeks of age. When given at this early age, a single dose protected infants for at least 2 months; two doses gave protection at least to the age of 18 months. After that, a booster can be given, but the Hopkins researchers believe that generally the child's immune system by then has matured sufficiently to fight the Hib organism.

"Early vaccination is vital to the success of Hib-OMPC," said Santosham.

"We proved that this is the only vaccine available today which can be safely administered to an infant only 6 weeks old," Santosham said, "and which provides protection beginning with the first dose."

The vaccine has been developed by Merck Sharp & Dohme, who supported the research.

A Multistate Outbreak of Hepatitis A Caused by the Consumption of Raw Oysters

Jean-Claude A. Desenclos, MD, Karl C. Klontz, MD, MPH, Michael H. Wilder, MD, MS, Omana V. Nainan, MS, Harold S. Margolis, MD, and Robert A. Gunn, MD, MPH

ABSTRACT

Background. **In August 1988 we investigated a multistate outbreak of hepatitis A caused by Panama City, Florida, raw oysters.**

Methods. **Cases of hepatitis A (HA) with onset in July–August 1988 were identified among persons who ate seafoods harvested in the coastal waters of Panama City, Florida. We conducted a case-control study, using eating companions of case-patients, and calculated attack rate (AR) per 1000 dozen raw oysters served. Enzyme immunoassay (EIA) and a polymerase chain reaction (PCR) technique were performed on samples of raw shellfish obtained from Panama City coastal waters.**

Results. **Sixty-one case-patients were identified in five states: Alabama (23), Georgia (18), Florida (18), Tennessee (1), and Hawaii (1). We found an increased risk of HA for raw oyster eaters (odds ratio = 24.0; 95% confidence interval = 5.4–215.0; $P < .001$). The AR of HA in seafood establishments was 1.9/1000 dozen raw oysters served. The EIA and PCR revealed HA virus antigen and nucleic acid in oysters from both unapproved and approved oyster beds, in confiscated illegally harvested oysters, and in scallops from an approved area.**

Conclusions. **The monitoring of coastal waters and the enforcement of shellfish harvesting regulations were not adequate to protect raw oyster consumers. More emphasis should be placed on increasing public awareness of health hazards associated with eating raw shellfish. (*Am J Public Health.* 1991;81:1268–1272)**

Introduction

The consumption of raw or undercooked shellfish has caused numerous outbreaks of bacterial and viral enteric diseases worldwide.[1–18] In the United States, outbreaks of hepatitis A (HA)[4–8] and of Norwalk gastroenteritis[15–17] caused by the consumption of raw shellfish have been reported regularly since 1962, and raw or undercooked clams or oysters were implicated as the most frequent vehicle of infection.[4–8,13–17] To prevent the occurrence of such outbreaks, many states have adopted regulations to prevent the harvesting and distribution of shellfish contaminated by human fecal pathogens.[19] However, these regulations have not prevented raw shellfish–borne outbreaks of HA[8] and Norwalk virus gastroenteritis.[16,17] These outbreaks were linked to shellfish by epidemiologic data; however, to our knowledge hepatitis A virus (HAV) has not been identified in implicated shellfish during outbreaks in the United States.

We describe a large multistate outbreak of HA caused by the consumption of Florida raw oysters. During this outbreak we were able to identify HAV in illegally harvested oysters that we believe were the source of infection, using an enzyme immunoassay (EIA) and in vitro nucleic acid amplification, a method potentially applicable to monitoring the microbiological suitability of shellfish intended for human consumption.

Methods

Background

In early August 1988, cases of HA were reported among Alabama patrons of Panama City, Florida, seafood restaurants and oyster bars. Panama City, located on the Gulf of Mexico, is surrounded by three bays (East, West, and North Bay) where shellfish, particularly oysters, grow naturally. Registered fishermen harvest oysters according to a management plan monitored by the Florida Department of Natural Resources. Oyster beds are classified as approved, conditionally approved, or prohibited for harvesting according to factors such as rainfall, water salinity, and water coliform counts. For 7 years before 1988, the bays surrounding Panama City were closed to oyster harvesting each year from June 1 to August 31. However, in 1988, summer harvesting was permitted in North Bay.

Epidemiologic Investigation

A case of HA was defined as a person who was IgM antibody positive to HAV, or who was diagnosed with HA by a physician during July or August 1988, and

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Jean-Claude A. Desenclos and Robert A. Gunn are with the Division of Field Epidemiology, and Omana V. Nainan and Harold S. Margolis are with the Hepatitis Branch of the Centers for Disease Control, Public Health Service, Atlanta. Karl C. Klontz and Michael H. Wilder are with Disease Control, Florida Department of Health and Rehabilitative Services, Tallahassee, Fla.

Requests for reprints should be sent to Jean-Claude A. Desenclos, European Centre for the Epidemiological Monitoring of AIDS, Hôpita National de Saint-Maurice, 14 rue du Val d'Osne, 94410 Saint-Maurice, France.

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Note. The inclusion of trade names in this article is for identification only and does not imply endorsement by the Department of Health and Human Services or the US Public Health Service.

who had attended a seafood event 10–50 days before symptom onset. Cases were ascertained by contacting neighboring states, all local Florida health departments, and Panama City hospitals and physicians, and through interviews with case-patients and managers of seafood establishments.

We conducted a case-control study in August 1988. All case-patients reached by telephone were included in the case group. The non-ill meal companions named by each patient (0 to 3 companions named per patient) were included in the control group if they did not develop any symptom of HA during the 10- to 50-day period following the seafood meal, if they did not report a past history of HA or jaundice, and if they had not received immune globulin during the previous 3 months. Serum specimens were not obtained to detect asymptomatic infection among controls. A standardized questionnaire was used for telephone interviews with cases and controls to ascertain health status and food consumption during the seafood event.

The direct and indirect costs of the outbreak were estimated from medical information obtained during the case-control study. Direct cost was calculated using mean routine medical fees at local hospitals; indirect cost was calculated at a mean salary of $100 per day of work lost.

To estimate attack rates (ARs) for oyster bar patrons, we reviewed oyster invoices and records in Panama City seafood establishments where patients had eaten. ARs were calculated per 1000 dozen raw oysters served by week and by establishment.

Crude odds ratios (OR) and 95% exact confidence intervals (CI)[20] were calculated for each food and exposure investigated. Statistical tests used included χ^2 with Yates correction, the two-tailed Fisher exact test, the χ^2 for trend, and the Student *t* test for continuous variables. When case-patients had eaten in more than one restaurant during the exposure period, the analysis of data was restricted to the restaurant which had reported other case(s) being exposed on the same day.

Environmental Investigation

We traced the origin of oysters by reviewing oyster bag tag numbers recorded by seafood establishments. Information on illegal oyster harvesting activities in Panama City coastal waters was obtained from the Florida Marine Patrol. Coliform count records from shellfish growing areas and at a sewage treatment plant (plant A) were reviewed. We inspected Panama City approved and unapproved oyster growing areas to search for sources of fecal pollution and collected samples of oysters and water from approved (North Bay) and two unapproved areas (Watson and Beatty Bayou), scallops from a scallop growing area (Shell Island), and sludge from another sewage treatment plant (plant B) located in Watson Bayou. We also obtained frozen raw oysters that had been illegally harvested in Watson Bayou in early August 1988 and confiscated by the Florida Marine Patrol.

Laboratory Investigation

Coliform counts were done on water and oyster meat samples using the three tube five dilution technique.[21] For the detection of HAV antigen and nucleic acid, individual shellfish (oysters and scallops) were removed from the shell; the digestive tract, mantle, and muscle were minced; and material from two specimens was pooled, homogenized in phosphate buffer saline + 0.1M ethylene diamine tetra acetate (EDTA), and clarified by centrifugation. Separate instruments were used to prevent cross-contamination of specimens. Supernatants of the homogenated shellfish were tested for HAV antigen using a previously described EIA.[22] HAV was quantitated from a standard curve prepared from cell culture–derived HAV, strain HAS-15,[23] and diluted in oyster meat homogenate supernatant fluid from oysters not implicated in an illness outbreak. Sludge samples from the sewage treatment plant were prepared as a 10% (w/v) suspension, sonicated, and clarified by centrifugation, and the supernatant was tested for HAV antigen.

Specimens (resuspended homogenate pellets, homogenate supernatant, and sludge) were tested for HAV-RNA by in vitro nucleic acid amplification using the polymerase chain reaction (PCR). HAV-RNA was obtained by immunocapture of virus with the capture antibody used in the HAV-Ag EIA[24] followed by heating at 900°C in reverse transcriptase buffer for 3–5 minutes. The oligonucleotide primer complementary to HAV-RNA was annealed at 90°C for 3 minutes, cooled on ice, and cDNA was synthesized at 42°C for 1 hour with 200 μmoles dNTP's, 5 units AMV reverse transcriptase, and RNAscin.[25] The second primer was added and HAV-cDNA was converted to double-stranded DNA and amplified with a heat stable (Taq) polymerase in a thermal cycler for 30 cycles (DNA Thermal cycler, Perkin-Elmer Cetus, Norwalk, Conn). Amplified nucleic acid products were analyzed by electrophoresis at 10 volts/cm on a horizontal gel (Hoefer Scientific Instruments, San Francisco, Calif) in a composite agarose gel of 3% NuSieve and 1% Sekem (FMC Bioproducts, Rockland, Md) prepared in Tris-borate-EDTA buffer containing 0.5 μg/ml ethidium bromide. DNA specific for the amplified region of the HAV genome was identified by hybridization using a 32P kinase-labeled oligonucleotide probe that was internal to the two limiting primers used in the PCR.[25]

Results

Epidemiologic Investigation

Sixty-one case-patients were identified; all but two had been tested and were IgM anti-HAV positive. Of 53 patients interviewed, all but 1 stopped working while ill and 17 (32%) were hospitalized (range = 2–21 days, median = 4 days). The duration of illness ranged from 7 to 49 days (median = 21 days), and the median number of visits to a physician while ill was four per patient (range = 1–13 visits). There were no deaths; seven secondary cases occurred among household contacts of case-patients. The total cost of the outbreak was approximately $200 000: $73 000 for direct costs, i.e., medical care and the outbreak investigation, and $130 000 for indirect costs associated with lost work days.

Illness onset ranged from July 4 to August 15 (Figure 1), the exposure period from June 3 to July 21, and the incubation period from 16 to 48 days (median = 29 days). Case-patients lived in Alabama (23 persons), Georgia (18), Florida (18), Tennessee (1) and Hawaii (1). Case-patients ate in seafood restaurants or oyster bars in Panama City (53 persons, 86.9%), in a restaurant in Alabama that served oysters harvested in Panama City (3; 4.9%), and at private family gatherings in Panama City (5; 8.2%). Age of patients ranged from 8 to 60 years (median age = 31 years), and 49 (80.3%) were male.

Fifty-three case-patients and 64 controls were enrolled in the case-control study. Cases tended to be younger than controls (mean age = 27.5 and 32.2 years, respectively; $P = .02$) and were more likely to be male (81 vs 39%; $P < .001$). The risk of HA was greater for patrons who ate raw oysters than for patrons who did not (OR = 24.0; $P < .001$) (Table 1) and increased with the number of raw oysters ingested (Table 2). Of the two case-patients who did not eat raw oysters, one ate baked oysters, and the other ate raw

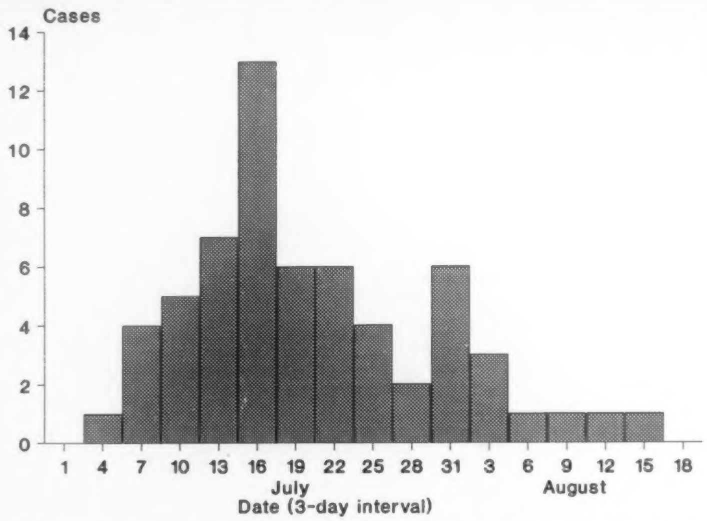

FIGURE 1—Cases of hepatitis A by date of onset, oyster associated hepatitis A outbreak, Florida, July–August 1988.

TABLE 1—Seafoods Consumed by Case-Patients and Controls, Oyster-Associated Hepatitis A Outbreak, Florida, July–August 1988

Seafood	Cases (n = 53) No.	Cases (n = 53) %	Controls (n = 64) No.	Controls (n = 64) %	Odds Ratio	95% CI[a]	P
Raw oysters	51	96	33	52	24.0	5.4–214.7	<.001
Fried or baked oysters	7	13	13	20	0.6	0.2–1.8	.4
Steamed oysters	1	2	5	8	0.2	0.01–2.1	.2
Raw shrimp	3	6	1	2	3.8	0.3–201.5	.3
Cooked shrimp	17	32	30	47	0.5	0.2–1.2	.1
Steamed clams	1	2	3	5	0.4	0.01–5.1	.6
Fish	17	32	18	28	1.2	0.5–2.9	.8

[a]CI = confidence interval.

TABLE 2—Number of Raw Oysters Eaten by Oyster-eating Case-patients and Controls, Oyster-Associated Hepatitis A Outbreak, Florida, July–August 1988

Number of Raw Oysters Eaten	Cases (n = 51) No.	Cases (n = 51) %	Controls (n = 33) No.	Controls (n = 33) %	Odds Ratio	95% CI[a]
1–6	6	12	18	55	1.0	—
7–12	20	39	11	33	5.5	1.5–21.5
>12	25	49	4	12	18.8	3.9–99.3

χ^2 for trend=20.0, $P < .001$.
[a]CI = confidence interval.

scallops caught at Shell Island beach. No other food items were associated with illness (Table 1), and the number of raw oysters eaten (>12 vs <12) had no effect on incubation period (<30 days; 55% vs 48%; rate ratio [RR] = 1.1; P = .9), duration of illness (>20 days; 79% vs 74%; RR = 1.1; P = .9), or hospitalization rate (29% vs 33%; RR = 1.1; P = .9).

During the exposure period, the 11 Panama City establishments associated with cases served approximately 27 900 dozen raw oysters. The median amount served per week was 3500 dozen (range = 1950–4800) (Figure 2), and the median number served by each establishment was 1994 dozen (range = 1050–8230). The overall AR was 1.9 cases per 1000 dozen raw oysters served with a peak during the third week of June (Figure 2). ARs by establishment ranged from 0.1 to 18 cases per 1000 dozen raw oysters, and most case-patients (44/52; 84.6%) had eaten raw oysters in one of four restaurants (A–D, Table 3).

Environmental Investigation

Seafood establishments and private individuals purchased raw oysters from wholesalers who obtained most of their oysters from Panama City–registered fishermen. Panama City bays met the defined safety standards at the time they were opened. Information gained from the Florida Marine Patrol indicated that illegal oyster harvesting at night was common during the summer of 1988, particularly in Watson Bayou where bootleggers arrested in early August had obtained multiple bags that they sold directly to some restaurants and oyster bars in Panama City.

Inspection of Watson Bayou showed failing septic tanks, boat sewage disposal, and sewage treatment plant sludge in proximity to unapproved oyster beds. Review of coliform counts at the effluent outfall of sewage treatment plant A was high [>16 000 most probable number (MPN)] from March through April 1988 and again in July, but not during May and June. This plant discharged approximately 1 mile from Watson Bayou.

Laboratory Results

Coliform counts higher than the maximum allowed were found in the water of two unapproved beds, although the oyster meat of only one of the beds was abnormal, and no human fecal pollution was detectable in the approved areas (Table 4). One of the three oyster pairs from approved beds of North Bay, all four scallops obtained from Shell Island, two oyster pairs from one of the beds sampled in Watson Bayou, and Watson Bayou oysters confiscated from a bootlegger were positive for HAV antigen by the EIA. HAV-RNA was detected by PCR in all oyster samples positive for HAV antigen by EIA (Table 4). However, two oyster samples HAV antigen negative by EIA were positive for HAV-RNA by PCR including a sample from Beatty Bayou, which was unapproved, and a sample from North Bay approved oyster beds (Table 4). Suspensions of sludge obtained from the sewage treatment plant were negative for HAV antigen and for HAV-RNA by PCR.

Discussion

This oyster-borne outbreak of HA is the largest reported in the United States since the 1973 Louisiana outbreak that involved 263 cases.[8] Although all case-patients recovered from their illness, the economic burden of the outbreak was high.

The risk of illness increased with the number of raw oysters consumed, suggesting a dose-response effect. However, there was no relationship between the quantity of raw oysters consumed and the incubation period, the duration, or the severity of illness. These findings suggest that the number of oysters eaten increased the chance of eating an infectious oyster rather than increasing the amount of virus ingested, which has been shown to decrease the incubation period.[26]

Although the case-control study showed that males had a greater risk of HA, this association is probably related to the use of eating companions of case-patients as controls: Florida adult males are more likely to eat raw oysters than females (Florida Behavioral Risk Factor Survey, Department of Health and Rehabilitative Services, unpublished data, 1988) and therefore to develop oyster-borne HA. Thus, controls as meal companions of case-patients were more likely to be females (spouse or girl friend).

Although the environmental investigation was not successful in definitively identifying the origin and the distribution of the contaminated oysters, we believe that contaminated oysters illegally harvested from unapproved waters were responsible for the vast majority of cases. However, we cannot exclude the possibility that some cases may have been caused by contaminated oysters harvested from approved beds of North Bay because both EIA and PCR identified HAV in oysters sampled there. Nevertheless, several factors strongly suggest that this outbreak was the result of bootlegging. First, because North Bay is the most common source of raw oysters served in Panama City, the observation of higher ARs of HA in four establishments suggests that more contaminated oysters were sold in those four establishments than in the other establishments. Second, according to anecdotal reports, many oysters had been illegally harvested from Watson Bayou in June and July 1988 and sold to specific seafood establishments in Panama City. Third, HAV was identified in oysters obtained directly from Watson Bayou and in oysters confiscated from a person who was arrested for illegal harvesting there. We also believe that the opening of Panama City coastal waters to summer harvesting after 7 years of summer closure facilitated the sale of the locally bootlegged oysters.

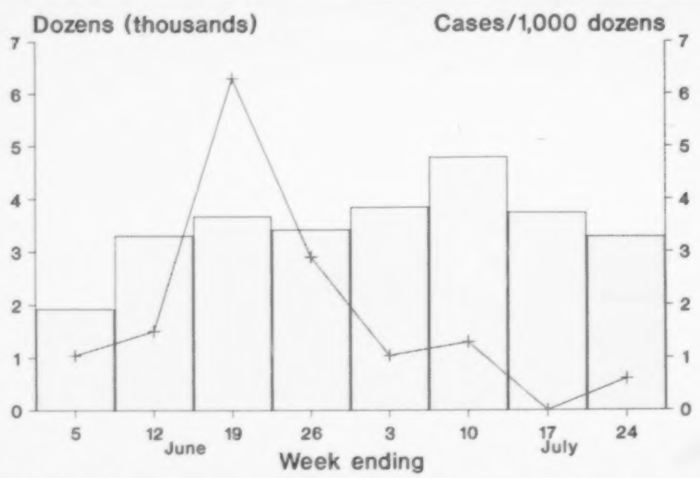

FIGURE 2—Raw oysters served (bar) and attack rate of hepatitis A per 1000 dozen raw oysters served (line) by week of exposure, oyster-associated hepatitis A outbreak, Florida, June–July 1988.

TABLE 3—Hepatitis A Attack Rate by Panama City Seafood Restaurants or Oyster-Bars[a] of Raw Oyster Consumption, Oyster-Associated Hepatitis A Outbreak, Florida, July–August 1988

Restaurant	No. of Cases	Cases per 1000 Dozen Raw Oysters
A	24	18.0
B	9	5.4
C	6	2.5
D	5	1.1
Others[b]	8	0.4
All	52	1.9

[a] n = 11.
[b] n = 7.

The raw scallops, eaten by one case-patient not exposed to oysters, were harvested off Shell Island where scallops were also found to contain HAV. This area is approximately 5 miles downstream from Watson Bayou, and a common source of fecal pollution may explain the contamination of shellfish from both locations.[14,15,18,27]

In the past, the isolation of viral pathogens from shellfish implicated in disease outbreaks has been difficult.[27] The finding of HAV by immunoassay in oysters obtained from the various harvesting areas indicated high levels of contamination. Even in the approved areas, oyster and scallop samples were positive for HAV by either or both EIA and PCR, which suggests that contamination may have been more widespread than was originally suspected. Virus isolation by cell culture was not attempted because of the extreme difficulty in growing wild-type HAV from environmental samples.[28] While direct proof of infectivity of either HAV-antigen or HAV-RNA positive oysters was not obtained, the epidemiologic evidence for infectivity was substantial, and it has been shown that HAV remains viable for long periods of time in water.[28]

The application of rapid viral diagnostic methods (EIA or PCR) for the identification of HAV and presumably other viral pathogens in shellfish has the potential for assuring the quality of this raw food product. However, these methods must be verified by infectivity studies in cell cul-

TABLE 4—Coliform Counts and Hepatitis A Virus Identification in Water and Shellfish Samples, Oyster-Associated Hepatitis A Outbreak, Florida, July–August 1988

Sample	Water Coliform[a]	Shellfish		
		Coliform[b]	HAV-EIA	HAV-PCR
Approved areas				
North Bay[c]	13	<20		
Pair 1			+	+
Pair 2			−	+
Pair 3			−	+
Shell Island[d]	NA	NA	+	+
Nonapproved areas				
Watson Bayou 1[c]	240	170	+	+
Watson Bayou 2[c]	NA	NA	−	−
Beatty Bayou[c]	30	330	−	+
Confiscated oysters	NA	NA	+	+

Note: HAV = hepatitis A virus; EIA = enzyme immunoassay; PCR = polymerase chain reaction; + = positive test result; − = negative test result; NA = not available.

[a] Most probable number per 100 cc; maximum level permitted for water is 14 per 100 cc.

[b] Most probable number per 100 g; maximum level for shellfish is 230/100 g.

[c] Oysters.

[d] Scallops.

ture or nonhuman primates. This outbreak points out the failures of current regulations in assuring the quality of raw shellfish because enforcement cannot effectively prevent illegal harvesting, and it suggests that even approved areas may be contaminated with HAV. Emphasis must be placed on increasing the public awareness of the health risks associated with the consumption of raw or undercooked shellfish. □

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Bleach Programs for Preventing AIDS among IV Drug Users: Modeling the Impact of HIV Prevalence

Joanna E. Siegel, ScD, Milton C. Weinstein, PhD, and Harvey V. Fineberg, MD, PhD

ABSTRACT

Background. **The growing importance of drug use as a mode of HIV transmission has led to increased attention to AIDS prevention among intravenous drug users (IVDUs). This analysis examines the effectiveness of bleach distribution, a program to prevent HIV transmission via shared needles.**

Methods. **We used a Markov model to assess the role of the initial HIV prevalence among drug users in determining the effectiveness of bleach programs. The model incorporates survey data on risk behaviors and published information describing HIV incubation and mortality. It predicts life expectancy for cohorts of IVDUs with and without a bleach program to estimate program effectiveness.**

Results. **We found that bleach programs can produce the greatest life-year savings in areas of low HIV prevalence. In the lowest prevalence scenario (0.02 initial prevalence), initiation of the program resulted in a projected savings of 2.3 life years per HIV-negative drug user, compared with 1.7 and 1.3 years under medium (0.25) and high (0.60) prevalence, respectively.**

Conclusions. **While bleach programs are beneficial in all groups of IVDUs, these results highlight the advantages of introducing bleach programs early, when prevalence is still comparatively low in a drug-user population. (*Am J Public Health.* 1991;81:1273–1279)**

Introduction

The proportion of AIDS cases occurring among persons whose primary risk factor is intravenous drug use has increased steadily since the beginning of the AIDS epidemic. In 1986, drug use was the primary risk factor for 16.8% of reported AIDS cases, compared to 21.5% currently.[1,2] Among new cases, the proportion is even higher.

Intravenous drug users (IVDUs) risk HIV infection from sexual contact and from sharing injection equipment—needles, syringes, or other items—with infected individuals. While protective changes in behavior have been reported among homosexuals, most IVDUs continue to place themselves at risk.[3,4] The growing importance of drug use as a mode of HIV transmission has increased attention to AIDS prevention in IVDU communities.

This analysis used a mathematical model to explore the effectiveness of one type of intervention, bleach distribution, for preventing AIDS among IVDUs. Bleach programs employ outreach workers to distribute small bottles of bleach, which IVDUs use to disinfect their injection equipment.[5] An important argument for the implementation of bleach programs in the United States has been that they are more politically feasible than needle exchange programs. However, arguments have been made for their use even where needle exchange programs are well established.[6]

Our study addressed the targeting of bleach programs, asking whether and to what extent initial HIV prevalence among IVDUs influences program effectiveness. A variety of obstacles, such as the difficulties of conducting longitudinal surveys of transient IVDU populations, prevent studies of existing programs from readily answering this question. Modeling therefore provides a practical means for studying the conditions that determine the success of bleach programs.

Methods

The model we used to predict bleach program effectiveness is a Markov (state-transition) model[7] similar to others developed to simulate the AIDS epidemic.[8-12] The model has four general states describing a hypothetical cohort of IVDUs (Figure 1). All IVDUs begin in an HIV-negative state. During each 1-year period, they face a risk of acquiring HIV infection from a needle-sharing partner and moving to HIV-positive status. Once infected, they face an annual probability of developing AIDS and then of dying from AIDS. HIV-negative and HIV-positive IVDUs die of other causes during the simulation. Death from non-AIDS causes is assumed to be negligible once an IVDU has developed AIDS.

We estimated probabilities of transition between model states from the literature, as described below. The model allows for recruitment of new HIV-negative drug users into the drug-using population. It also includes already HIV-positive

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The authors are with the Department of Health Policy and Management, Harvard School of Public Health, and the Center for Policy and Education, Harvard AIDS Institute, Boston.

Requests for reprints should be sent to Joanna E. Siegel, ScD, Department of Health Policy and Management, Harvard School of Public Health, 677 Huntington Avenue, Boston, MA 02115.

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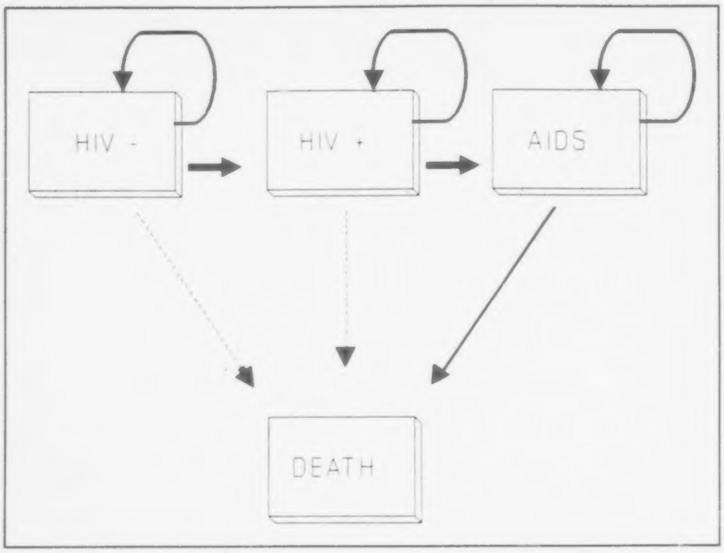

FIGURE 1—Markov Model of HIV transmission among Intravenous Drug Users.

IVDUs, who contribute to the prevalence of HIV in the community. Program effectiveness is calculated only for the initially HIV-negative IVDU cohort, however, and not for the community as a whole.

We use two measures of program effectiveness. The first, median infection-free survival, is the number of years until half of the modeled cohort become infected. The second is life expectancy, the average survival of cohort members.[7] Both measures capture the protective effect of bleach use on the individual IVDU as well as the reduction in HIV prevalence which protects other cohort members. Both also reflect continuing exposure to HIV—the fact that an IVDU who escapes infection for 1 year will be exposed again the following year. This long-term risk significantly increases lifetime probabilities of infection and is an important determinant of the effectiveness of interventions in an IVDU population.

Median infection-free survival and life expectancy are predicted both without and with a bleach program; effectiveness is defined as the improvement in these measures resulting from the program. We compared predicted effectiveness assuming programs are imple[illegible] where the HIV prevalence among drug users is low, medium, and high.

The analysis was conducted in three steps described in the results section: validation, baseline analysis, and sensitivity analysis. Change in life expectancy rather than median infection-free survival is used in the sensitivity analysis because its calculation is more precise. Computing is done on an IBM personal computer using the SMLTREE program (James P. Hollenberg, Copyright © 1985).

This model does not consider sexual transmission either within or outside of the drug-using population. While sexual transmission between individuals who also share needles is probably a secondary source of infection, sexual transmission to nondrug users is an important policy concern not directly addressed by our analysis.

Transition Probabilities

Risk of HIV infection. The cumulative annual probability of infection for IVDUs in the model depends on both personal behaviors and factors external to the IVDU (Appendix A). The risk behaviors include number of needle-sharing partners (m), injection frequency (i), and frequency of needle-sharing (s). The HIV prevalence among drug users in the community determines the probability of selecting an infected injection partner (p). A final component of risk is the probability (e) that the virus is transmitted when an IVDU shares injection equipment with an infected person. This probability is decreased when IVDUs disinfect their equipment with bleach (u,k). The model assumes that the risk of infection from each shared-needle contact with a given partner is constant and independent of previous contacts.

We used data from the Urban Health Study, a survey of IVDUs in San Francisco, to estimate drug-use behaviors.[13] From a sample of 603 interviewed in 1988, we selected those who reported both injecting drugs and sharing equipment during the previous year (n=251). The IVDUs we modeled were thus an active and at-risk subgroup of drug users.

Epidemic models suggest that heterogeneity of risk behavior may have an important effect on the course of the AIDS epidemic.[14,15] We considered heterogeneity among drug users by defining three separate groups at risk of infection. The groups are distinguished by frequency of injection (i), a behavior found to be associated with HIV infection, as our model specifies (Appendix A).[16–18] We used a partitioning algorithm, which minimizes total within-group variation, to differentiate groups within the Urban Health Study data. The resulting groups included a large number of "low-injector" IVDUs (≈1 injection per day); a "medium" group (≈4 injections per day); and a small "high" group (≈8 injections per day) (Appendix B). We determined the average proportion of injections shared (s) and number of needle-sharing partners (m) for each subgroup directly from the data.

Because injection frequency is associated with HIV infection, we assumed that the IVDU subgroups have different HIV prevalences prior to the bleach intervention (Appendix B).[16] In the course of the simulation, prevalence continued to be determined separately for the three groups. IVDUs in the model thus face a different annual probability of infection depending on their injection practices.

We estimated the probability of viral transmission per infected needle contact (e) based on the transmission in needlestick accidents among health care workers. Combining data from a number of studies, Friedland and Klein estimate a risk between 0.0013 and 0.0039 (upper confidence level = 0.0076).[19] We used 0.003 in our baseline analysis and, because this estimate is quite uncertain, a wide range was tested in the sensitivity analysis.

Bleach programs are actually multiservice outreach programs, referring IVDUs to treatment and providing information on risk behaviors. While other advantages of bleach programs may pertain, our model considers only the direct impact of bleach use on survival. We assume bleach effectiveness (k) of 85%—despite the virtually complete inactivation of HIV shown in laboratory studies of bleach[20-22]—to reflect problems in actual usage. If IVDUs disinfect only their needles, for example, the virus may be transmitted by other shared equipment. The probability that the IVDU uses bleach at each injection (u) is based on studies of IVDUs in San Francisco: 0.10 before the implementation of a program and 0.55 after.[13,23,24]

Risk of developing AIDS. Our estimates of the incubation period (time from HIV infection until the onset of AIDS) are based on data from the San Francisco Clinic Cohort, a group of HIV-positive homosexual men, because no similar studies document incubation in IVDUs.[25,26] We derive annual probabilities of developing AIDS from a Weibull model fit to these data, as described by De Gruttola and colleagues.[27,28] The Weibull model allows the estimated annual risk of developing AIDS, given HIV infection, to increase with time. The Markov model is adapted to reflect these time-dependent probabilities.

Annual probability of death. Probabilities of death for IVDUs with AIDS are based on a study of 1- to 5-year survival by Rothenberg et al.[29] Advances in treatment may ultimately increase survival, although the recent improvements among homosexual men have not been documented among IVDUs.[30] The impact of future treatments will depend on whether longer survival is associated with extended drug use. Our model currently specifies that persons with AIDS transmit HIV for only 1 year. Our conclusions would not be affected by a longer duration of infectivity, because most of the infection in the cohort takes place before AIDS develops. However, predicted life expectancy would increase. For this reason, our estimates of life expectancy are appropriate only for comparative purposes within this analysis.

Age-specific mortality from non-AIDS causes is derived from the US life tables for males[31] and adjusted to reflect higher risks of death from overdose and violence associated with drug use.[32] Not included in these probabilities is a substantial increase in mortality coincident with the AIDS epidemic, believed to be HIV-related.[33] The effect of an additional component of excess mortality among HIV-positive IVDUs is tested in the sensitivity analyses.

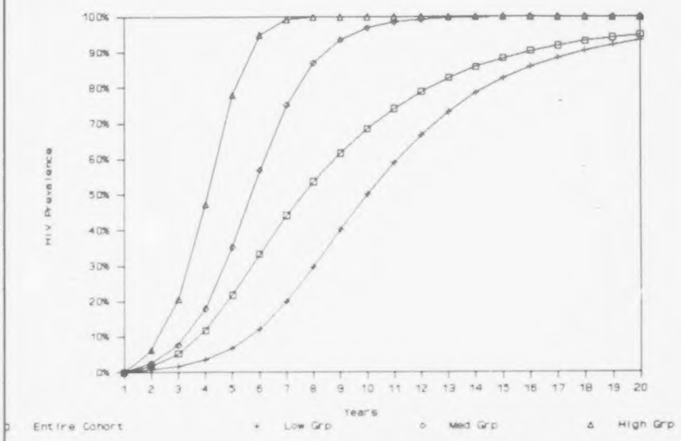

FIGURE 2—Increase in prevalence by risk group (initial prevalence 0.02).

Results

Validation: Model Dynamics and Predictions of Life Expectancy

One indicator of the validity of a model is that it obtains expected results under simple conditions. In this section, we examine how simulated life expectancy responds to variations and to extreme values of model parameters.

In the absence of AIDS, the model predicts an average undiscounted life expectancy of 26.64 years for a 35-year-old IVDU. For an IVDU who becomes infected with HIV at age 35, life expectancy is only 9.58 years. Life expectancy in the modeled cohort drops sharply when the virus is introduced, reflecting the rapid transmission of even low levels of HIV in this high-risk population.

The speed of dissemination of HIV in a drug-injecting population is illustrated by the increase in annual prevalence over time (Figure 2). When the initial HIV prevalence in the community is 2%, for example, cohort prevalence reaches 12% after 3 years and 53% after 7 years in the absence of an intervention program. (Community prevalence, which includes recruited IVDUs, is slightly lower.) This pattern of increasing annual prevalence falls within the range of trends observed in cities including New York, Edinburgh, and several Italian cities.[34-36] Prevalence increases are predictably slower among IVDUs who inject relatively infrequently and more rapid among frequent injectors.

Risk factors such as the number of injections per year (i) directly affect predicted life expectancy. Life expectancy decreases slightly at low levels of injection (Figure 3) and then more rapidly as infection becomes more probable. At high injection frequency, IVDUs are infected rapidly, and more frequent injection does not significantly worsen the risk.

While the results of these simulations are difficult to validate empirically, the model demonstrates plausible responses to changes in parameter values. Of note is the very rapid fall in life expectancy that occurs in response to modest increases in determinants of risk, such as HIV prevalence, injection frequency, or probability of viral transmission.

Baseline Estimates of Bleach Program Effectiveness

The baseline analysis examines the influence of HIV prevalence on predicted bleach program effectiveness, using our best estimates for parameter values. We compared effectiveness at three levels of overall initial HIV prevalence: low (0.02), medium (0.25), and high (0.60). These levels correspond to 1987 estimates of HIV prevalence among IVDUs in Los Angeles, Boston, and New York, respectively.[37,38]

The baseline simulations indicate that a bleach program would slow the epidemic most successfully in a low-prevalence area (Table 1). The median infection-free survival increases from 5.9 years

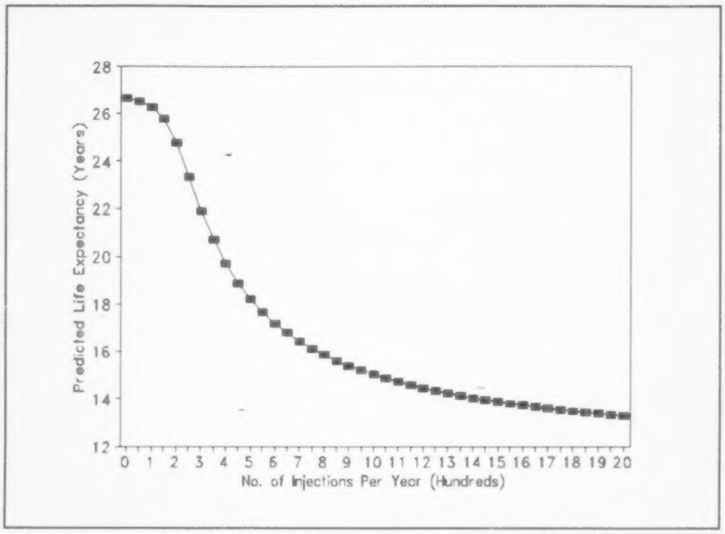

FIGURE 3—Predicted life expectancy by frequency of injection.

TABLE 1—Baseline Results for Low, Medium, and High Levels of Initial HIV Prevalence

Level	Initial Overall Prevalence[b]	Before Bleach Program	With Bleach Program	Change
Median infection-free survival, y[a]				
Low	0.02	5.9	8.6	2.7
Medium	0.25	2.9	4.4	1.5
High	0.60	1.8	2.9	1.1
Life expectancy, y				
Low	0.02	15.54	17.86	2.32
Medium	0.25	13.43	15.11	1.68
High	0.60	12.62	13.94	1.32

[a] Number of years elapsing before 50% of cohort is HIV-infected.
[b] Initial prevalence in population of cohorts and recruits; weighted average of subgroup prevalences.

TABLE 2—Gain in Life Expectancy Following Program Implementation: Sensitivity Analysis on Probability of Bleach Use

Initial Overall Prevalence	Probability of Bleach Use After Program Implemented			Complete Bleach Use and Bleach Effectiveness[b]
	55%[a]	90%	100%	
0.02	2.32	6.72	8.76	11.11
0.25	1.68	5.47	7.65	13.22
0.60	1.32	4.53	6.59	14.02

[a] Baseline value.
[b] At bleach use 100% and bleach effectiveness 100%, no further HIV infection occurs. This level of success is unlikely but illustrates the case when the greatest gain occurs in the high-prevalence scenario, where life expectancy was lowest prior to the program.

to 8.6 years. The bleach program also slows the spread of HIV in medium- and high-prevalence areas, but the change is less pronounced.

Predicted life expectancy demonstrates the same pattern of effectiveness. The bleach program adds more years to the life of an HIV-negative IVDU in a low-prevalence area than in a high-prevalence area. In the lowest-prevalence scenario, the initiation of the program results in a projected savings of 2.3 life years per HIV-negative IVDU (Table 1). The savings is 1.7 in the medium-prevalence area and 1.3 years under high prevalence. The program is thus predicted to be more effective in areas where drug users are less likely to be exposed to the virus.

Sensitivity of Results

Sensitivity analyses determine whether varying the conditions in an analysis will alter its conclusion. Changes in our model have the general effect of worsening or improving the predicted course of the AIDS epidemic. The question is therefore whether plausible changes in the predicted epidemic course would modify our basic conclusion, i.e., that bleach programs are more effective in low-prevalence areas.

Bleach use and effectiveness. Program effectiveness remains greatest in the low-prevalence scenario at any level of bleach use (u) (Table 2). This occurs because bleach effectiveness (k) is less than perfect, so that HIV infection continues even if IVDUs use bleach at all injections. Transmission is greatly slowed when both bleach use and bleach effectiveness are assumed to be higher. For example, with bleach used at 90% of injections and 95% effectiveness, median infection-free survival increases from 5.9 years without the program to over 20 years. However, combined bleach use and bleach effectiveness well over 90% would be required to change our baseline results.

Interactions among IVDUs. We examined annual recruitment of new drug users up to 10% of the original cohort size, the equivalent of, for example, an estimated 20 000 new IVDUs in New York City each year. (Table 3). Although the addition of larger numbers of uninfected IVDUs dampens the yearly increase in prevalence, higher program effectiveness is still found in the lowest prevalence area.

Sensitivity analysis on patterns of mixing among IVDU subgroups obtains a similar result. In the baseline analysis, we assumed a variant of "preferred mixing," in which IVDUs share most but not all injections within their own subgroup (Appendix B).[39] Sensitivity analyses test the effect of restricted mixing (sharing only with others in the same group) and proportional mixing (sharing unrelated to the user's own group). Effectiveness is generally higher with restricted mixing because the large low-injection subgroup is isolated from higher-prevalence IVDUs. However, overall effectiveness remains greatest at low prevalence regardless of the mixing assumption.

Components of risk: Probability of viral transmission (e) and injection frequency (i). In the sensitivity analysis on the per-injection probability of HIV transmission (e), predicted effectiveness remains highest in low-prevalence areas. However, at the lowest values of e (0.001), predicted effectiveness in low- and medium-prevalence areas is nearly the same. Effectiveness in this range of e has begun to diminish at the lowest prevalence level and would soon fall below that in medium- and high-prevalence areas. Therefore, while the baseline conclusion does not change in the sensitivity analysis, the illustration in Figure 4 reveals how still lower values of this parameter would change our results.

The sensitivity analysis on injection frequency also examines a lower risk scenario. Lower risk groups of injectors are defined by assuming five daily injections for any drug user reporting more than that number. The three risk groups are then recompiled using the method described previously. In this analysis, effectiveness increases at all prevalence areas and remains highest in the low-prevalence area.

Mortality. Although increased mortality from non-AIDS causes decreases life expectancy in our model, the pattern of program effectiveness is not changed. Effectiveness increases somewhat at all prevalence levels and remains highest at low prevalence.

Discussion

Under our best assumptions about the AIDS epidemic among drug users, we predicted that bleach programs would be most effective in low HIV-prevalence areas. Bleach programs thus save more lives when IVDUs are less likely to be exposed to HIV. Intuitively, this occurs because IVDUs are more likely to avoid (or delay) infection in low-prevalence areas; they therefore gain more life years if bleach prevents infection at a particular point in time.

Sensitivity analyses suggest that our results are reasonably stable. Predicted effectiveness is highest in lower-prevalence areas, with two exceptions. If bleach use entirely or almost entirely eliminates HIV transmission, the program is most effective where loss of life was initially greatest—the higher-prevalence areas. Or, if HIV infection is initially extremely unlikely in the low-prevalence area, effectiveness becomes quite low because so few lives are endangered by the epidemic that bleach has few to save.

These are extreme cases, however. Generally, factors slowing the AIDS epidemic would increase, not decrease, the relative effectiveness of bleach distribution in low-prevalence areas. If IVDUs use bleach consistently and reduce risk in other ways, the combination will offer im-

TABLE 3—Gain in Life Expectancy Following Program Implementation: Sensitivity Analysis on Recruitment, Mixing, Viral Transmission, and Injection Frequency

Factors	Initial Overall Prevalence 0.02	0.25	0.60
Recruitment[a]			
No new entrants	2.11	1.52	1.22
Recruitment 5%[b]	2.32	1.68	1.32
Recruitment 10%	2.41	1.78	1.38
Mixing[c]			
Restricted	3.04	2.31	1.55
Preferred[b]	2.32	1.68	1.32
Proportionate	2.36	1.51	1.19
Probability of viral transmission per infected needle contact			
0.001	2.74	2.69	2.45
0.002	2.82	2.24	1.82
0.003[b]	2.32	1.68	1.32
0.004	1.89	1.32	1.01
0.005	1.59	1.08	0.82
Injection frequency			
Baseline[b] (max. 10/day)	2.32	1.68	1.32
Lower (max. 5/day)	2.72	1.83	1.45

[a] Recruitment is number of new HIV-negative IVDUs entering model each year as percentage of number in original HIV-negative cohort.

[b] Assumption in baseline analysis.

[c] Restricted mixing is sharing equipment only with others in the same injection-frequency group. In preferred mixing, medium- and high-injector groups share 15% of injections with a lower-risk group and remaining injections within their own group. Proportional mixing is sharing unrelated to the user's own group.

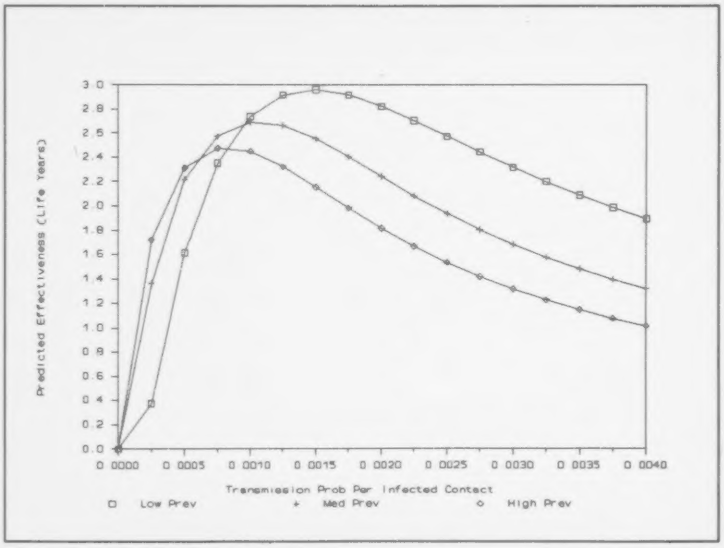

FIGURE 4—Predicted effectiveness by initial prevalence for assumptions of probability of HIV transmission per infected needle contact.

pressive life-saving potential, especially in low-prevalence areas (Table 2).

In the face of a dramatic and devastating illness such as AIDS, a reasonable response on the part of policymakers is to target resources to visibly distressed locations. What we wish to emphasize is the critical importance of preventive programs where AIDS may be less visible. It is in low-prevalence areas, before widespread infection has occurred, that bleach programs have their greatest potential for saving lives.

Several limitations should be considered before generalizing from the results of this model. There are limitations in the data, due to the intrinsic difficulty of surveying IVDUs.[40] Geographical differences in drug-use practices will affect generalizability. Some potential sources of heterogeneity have not been explored, for example, possible changes in viral infectivity over time. Our model does not reflect interruptions to drug use for periods of incarceration or drug treatment, which would lower the lifetime risk of contracting AIDS. These omissions will influence our predictions of life expectancy and may cause us to understate program benefits; they are less likely to affect our comparisons of program effectiveness across prevalence levels. As noted earlier, our analysis also excludes sexual transmission of HIV, a source of risk for IVDUs as well as their sexual partners and children. Nor are side-benefits of bleach distribution considered, such as education and referral to drug treatment.

In this analysis we have not assessed the cost or cost-effectiveness of programs in areas of differing HIV prevalence—and indeed the probable result of such an analysis is not obvious. However, because HIV prevalence among drug users can increase rapidly, our current opportunities for early intervention may be short-lived. The establishment of bleach programs in low-prevalence IVDU communities should become an explicit priority in a national effort to prevent AIDS. ☐

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APPENDIX A—Risk of Infection for HIV-Negative Intravenous Drug Users[a]

$$r_j = 1 - [p((1-s_j e(1-uk))^{i_j/m_j} + (1-p)]^{m_j}$$

r_j = cumulative annual probability of infection
j = group of IVDUs
p = probability of selecting an infected injection partner (a function of prevalence in the several risk groups)[b]
s_j = probability that needle is shared; i.e., probability that someone used the needle immediately prior to this injection
e = efficiency of transmission of HIV by infected needle, per injection
u = probability that bleach is used at each injection[a]
k = effectiveness of bleach in eliminating HIV virus from a needle and syringe
m_j = number of partners per year
i_j = number of injections per year

[a] Further discussion in references 41–43.
[b] These parameters are group-specific in some model versions.

APPENDIX B—Model Specifications for Baseline and Sensitivity Analyses

	Baseline Analysis[a]			Sensitivity Analysis[b]		
Specification	IVDU			Subgroup		
Group-specific parameters	*Low*	*Medium*	*High*	*Low*	*Medium*	*High*
j Proportion of cohort in subgroup	0.60	0.31	0.09	0.43	0.35	0.22
i Number of injections per year	337	1311	3018	212	841	1723
s Fraction of injections shared	0.35	0.26	0.27	0.36	0.32	0.22
m Number of needle-sharing partners	11	13	12	10	14	11
p HIV prevalence among IVDUs						
Overall prevalence 0.02	0.012	0.031	0.031[c]	—	—	—
Overall prevalence 0.25	0.169	0.346	0.346[c]	—	—	—
Overall prevalence 0.60	0.478	0.704	0.704[c]	—	—	—
Cohort parameters						
u Per injection probability of using bleach						
Before bleach program implementation	0.10			—		
After bleach program implementation	0.55			0.56–1.00		
k Effectiveness of bleach	0.85			0.55–1.00		
e Probability of viral transmission per infected needle contact	0.003			0.001–0.005		
Other model features						
Mixing pattern among subgroups	Preferred[d]			Restricted; proportionate[e]		
Recruitment of new IVDUs into population (fraction of initial cohort population)	0.05			0.00–0.10		

[a] Values for group-specific parameters obtained from Urban Health Study data (see text).
[b] Sensitivity analysis on group-specific parameters assumes maximum of 1825 injections per year. Subgroups are redefined from data to obtain values of j, i, s, and m for sensitivity analysis (see text).
[c] Prevalences for medium- and high-injector subgroups assumed to be the same
[d]. Medium- and high-injector groups share 15% of injections with lower risk group and remaining injections within their own group.
[e] Restricted mixing is sharing equipment only with others in same injection-frequency group; proportional mixing is sharing unrelated to user's own group.

Inpatient Morbidity among HIV-Infected Male Soldiers prior to Their Diagnosis of HIV Infection

ABSTRACT

Background. A natural history study of human immunodeficiency virus (HIV) disease was carried out among 1575 HIV-infected US Army men and 6220 demographically similar uninfected soldiers. Inpatient morbidity occurring up to 8 years prior to the date of HIV infection diagnosis among those men who became HIV infected was evaluated for both groups.

Methods. Incidence density rates were calculated for hospital admissions. Poisson regression was used to assess the trend in hospital admissions among those subsequently diagnosed with HIV infection. Prevalence ratios for discharge diagnoses were also calculated.

Results. Sixteen diagnoses/diagnosis categories occurred statistically more frequently among subsequently HIV diagnosed individuals than among those who remained uninfected. Among these were hepatitis B and abscess of anal/rectal region (6 to 8 years prior to HIV infection diagnosis); unspecified viral infection, enlarged lymph nodes, syphilis (3 to 5 years prior to HIV infection diagnosis); and diagnoses suggestive of acute retroviral syndrome (1 to 2 years prior to HIV infection diagnosis).

Conclusions. Data such as these may provide useful information to HIV surveillance efforts regarding patterns of morbidity experienced prior to HIV infection as well as to health care providers regarding patients at high risk for becoming infected with HIV. (*Am J Public Health.* 1991;81:1280–1284)

Philip O. Renzullo, MPH, John G. McNeil, MD, Lytt I. Gardner, PhD, and John F. Brundage, MD

Introduction

A natural history study of a large cohort of human immunodeficiency virus (HIV)-infected male soldiers and demographically similar uninfected soldiers was performed to investigate patterns of morbidity prior to and concurrent with early HIV infection. Hospital admission data during the period from HIV-infection diagnosis to 8 years prior to HIV diagnosis were analyzed to determine if the rate of hospital admissions differed for those later diagnosed with HIV infection compared to those who remained uninfected and if overall differences existed in the type and frequency of morbidity based on subsequent HIV-infection status.

Characterization of morbidity occurring around the time of HIV seroconversion is important for several reasons: (1) these signs and symptoms were not heretofore considered part of the clinical spectrum of HIV disease because they occur early in the clinical course, in some cases even before an individual suspects his or her HIV infection; (2) this information can be used by health care providers to assist in intervening early in the disease process; (3) it provides insight into the pathophysiologic processes that occur during the early stages of disease; and (4) it aids assessment of the burden and spectrum of morbidity that might be prevented by effective infection prevention programs.

Since October 1985, all military personnel on active duty and applicants for military service have undergone routine testing for HIV antibody (HIV-Ab).[1–4] Since the start of the screening program, the US Army HIV Data System (USAHDS) has archived the results of over 6 million HIV-Ab tests, approximately 2 million of them from active duty soldiers. The USAHDS maintains identifier-linked data files reflecting the dates and results of all HIV-Ab tests.

The US Army Medical Department also maintains a hospital discharge surveillance system called the Individual Patient Data System (IPDS). At the time of every discharge from an Army hospital, an entry is placed on IPDS consisting of demographic information, the primary discharge diagnosis and up to seven additional diagnoses, surgical procedures performed, and other medical care information. Thus, databases such as IPDS can be used to track hospitalized morbidity and medical care utilization over time for all beneficiaries of the Army health care system.

In addition to its ability to track morbidity among hospitalized cohorts, IPDS can be linked to other databases to track specific subcohorts of individuals. For example, when linked to the USAHDS registry of HIV seropositive individuals, IPDS provides a unique opportunity to evaluate the type and extent of inpatient morbidity in an otherwise healthy, employed HIV-infected population over an

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The authors are with the Department of Epidemiology, Division of Preventive Medicine, Walter Reed Army Institute of Research, Washington, DC.

Requests for reprints should be sent to Philip O. Renzullo, Department of Epidemiology, Division of Preventive Medicine, Walter Reed Army Institute of Research, Washington, DC 20307-5100.

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Note. The opinions or assertions contained in this paper are the private views of the authors and are not to be construed as official or as reflecting the views of the Department of the Army or the Department of Defense.

extended period both proximate to and remote from the time of HIV-infection diagnosis.

TABLE 1—Hospital Admissions, Person-Years, and Unadjusted Hospital Admission Incidence Density Ratios (cIDR) in Each of 8 Years Prior to HIV-Infection Diagnosis for HIV-Ab+ and Demographically Similar HIV-Ab− Individuals

Year Relative to Reference Date	No. Hospital Admissions		No. Person-Years		cIDR (95% CI)[a]
	HIV-Ab+	HIV-Ab−	HIV-Ab+	HIV-Ab−	
−8	64	143	623	2487	1.79 (1.33–2.39)
−7	65	187	707	2830	1.39 (1.05–1.84)
−6	74	222	811	3276	1.35 (1.03–1.75)
−5	94	224	927	3731	1.69 (1.33–2.15)
−4	117	263	1063	4258	1.78 (1.43–2.21)
−3	152	330	1233	5004	1.87 (1.54–2.26)
−2	194	381	1422	5744	2.06 (1.73–2.44)
−1	269	535	1575	6220	1.98 (1.72–2.29)

[a] CI = confidence interval.

Methods

HIV-infected male soldiers on the USAHDS registry with HIV-infection diagnosis dates between 1986 and 1988 were included in the study. The comparison group was composed of uninfected male soldiers identified from US Army personnel data bases. Four uninfected soldiers were matched to each infected soldier on the basis of age (±2 years), race (White, non-Hispanic; Black, non-Hispanic; Hispanic; Asian; other), rank (junior enlisted; senior enlisted; warrant officer; commissioned officer), and length of military service (±5 years). The file of study subjects was linked by social security number to the IPDS for retrieval of ICD-9-CM coded discharge diagnoses. A reference date (RD) was defined for each HIV-infected individual as the earlier of the date of his first confirmed positive HIV-Ab test or the date of any hospital admission specifically coded for AIDS or HIV infection. Uninfected soldiers were assigned the same RD as the HIV-infected soldier to whom they were matched.

Hospital admissions for each individual were allocated to the year in which they occurred relative to the RD. Due to the small numbers of admissions occurring more than 8 years prior to the RD, only those that occurred from 1 to 8 years prior to the RD were used in the analysis. To eliminate bias resulting from individuals with many hospitalizations, one admission per individual per year was used in the analysis. Similarly, duplicate diagnoses in the same year were eliminated, resulting in the occurrence of any specific diagnosis only once per person per year.

Statistical Analysis

For this analysis, HIV-Ab status was considered to be a marker for past exposures that resulted in hospitalizations. Because we were interested in determining whether the hospitalization pattern of HIV-infected individuals differed from the pattern of individuals not infected prior to the infected individuals being diagnosed with HIV, we selected the outcomes of interest to be hospital admissions and unique discharge diagnoses/diagnosis types.

Hospital admission incidence rates were calculated as time-specific incidence densities (ID) and defined as the number of new admissions per 100 person-years available in each year prior to the RD. The measure of association was the incidence density ratio (IDR) calculated by the formula $IDR = ID_1/ID_2$, where subscripts indicate demographic and time-specific strata. The Poisson method and a completely categorical model allowed for the calculation of the number of hospital admissions per stratum weighted by the follow-up time (p^*y) in each stratum, adjusted for age, race, and marital status.[5] Model fits were performed with the EGRET statistical package.[6] Ninety-five percent confidence intervals were calculated using the standard errors of β_k. The temporal trend associated with hospital admission incidence density ratios was determined over the 8 years prior to HIV infection diagnosis. The linear trend statistic calculated was the Poisson regression coefficient.

Prevalences (P) for unique discharge diagnoses, defined as the number of unique diagnoses of a particular type divided by the total number of diagnoses for both the HIV-infected and uninfected groups, were calculated over the combined 8-year period. Prevalence ratios (PR) for each unique discharge diagnosis defined as $PR_1 = P_{1(HIV\text{-}Ab+)}/P_{1(HIV\text{-}Ab-)}$, where subscripts indicate unique diagnosis types, were calculated for the combined 8-year period. Ninety-five percent confidence intervals were calculated using Taylor Series approximations.[7]

Results

A total of 1575 HIV-infected soldiers and 6220 uninfected soldiers were included in the study. The mean age for all subjects was 28 years. Thirty-four percent of subjects were White, non-Hispanic; 54% were Black, non-Hispanic; and 7% were Hispanic.

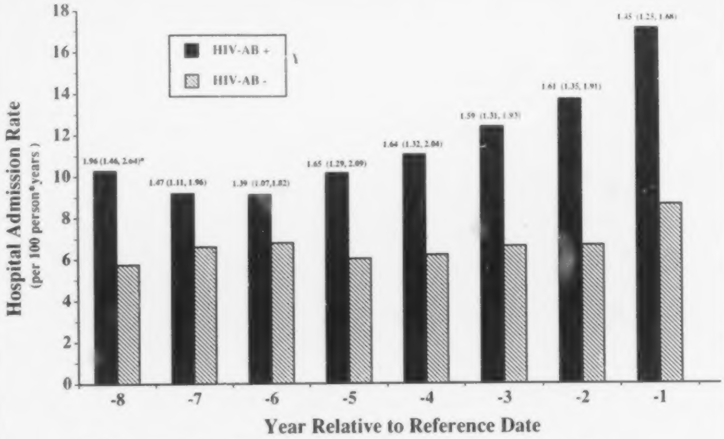
Hospital Admissions

There were 4602 hospital admissions between the RD and 8 years prior for all individuals in the study, 1469 among HIV-Ab positives and 3133 among HIV-Ab negatives. After removal of duplicate admissions within each year interval, infected soldiers experienced 1029 hospital admissions while uninfected soldiers experienced 2285 hospital admissions. Table 1 presents the number of hospital admissions and available person-years for HIV-Ab positive and HIV-Ab negative individuals in each year prior to the RD.

Figure 1 presents hospital admission rates for both groups over the 8 years prior to the RD. In general, admission rates for HIV-infected soldiers displayed a monotonic increase beginning 6 years prior to HIV-infection diagnosis. In contrast, admission rates among uninfected soldiers were nearly invariant. In comparison to the combined rates during the period 6 to 8 years prior to the RD, annual hospital admission rates increased overall by 79% in HIV-infected and 34% in uninfected soldiers. Admission rates were significantly greater in all years among those who were subsequently diagnosed with HIV infection. Further, the excess in hospital admissions among those with HIV infection became greater as the RD was approached (regression coefficient t test = -0.071, $P < .001$).

Individual Diagnoses

There were 1903 and 3671 diagnoses among HIV-infected and uninfected men,



* Incidence Density Ratio (95% Confidence Interval) adjusted for age, race, and marital status

FIGURE 1—Hospital admission incidence density rates and adjusted ratios for each of 8 years prior to the reference date.

TABLE 2—Diagnoses with Prevalence Ratios (PR) ≥2.0 that Occur Four or More Times over the 8-Year Follow-up Period among Individuals Subsequently Diagnosed with HIV Infection

Diagnosis	No. Diagnoses HIV-Ab+	No. Diagnoses HIV-Ab−	PR	95% Confidence Interval
Syphilis[a]	42	2	40.51	9.82–167.17
Enlargement of lymph nodes	49	3	31.51	9.33–100.95
Hepatitis B[b]	37	7	10.19	4.55–22.83
Unspecified hepatitis[c]	24	11	4.21	2.07–8.57
Circumcision	5	1	9.64	1.13–82.50
Hematologic disorders[d]	24	5	9.26	3.24–24.23
Adverse drug effect[e]	10	3	6.43	1.77–23.34
Urethral scarring requiring surgery (etiology unknown)	6	2	5.79	1.17–28.65
Acute bronchitis	6	2	5.79	1.17–28.65
Sexually transmitted diseases (not including syphilis)[f]	39	16	4.70	2.63–8.39
Fever of unknown origin	9	4	4.34	1.34–14.08
Chronic tonsillitis	11	6	3.54	1.31–9.55
Unspecified viral infection	58	50	2.24	1.54–3.25
Disorders of anal/rectal region[g]	46	40	2.21	1.46–3.38
Unspecified disorders of anus/rectum/GI tract[h]	30	27	2.14	1.28–3.59
Acute pharyngitis	21	19	2.13	1.15–3.95
Herpes zoster	4	1	7.72	0.86–68.99
Hepatitis A	5	2	4.82	0.94–24.83
Viral meningitis	4	3	2.57	0.57–11.48
Syncope/collapse	8	7	2.20	0.80–6.07
Mononucleosis	9	8	2.17	0.84–5.61

[a] Latent syphilis, secondary syphilis, unspecified syphilis, neurosyphilis.

[b] Viral hepatitis B (with or without hepatic coma), chronic hepatitis, other specified viral hepatitis.

[c] Unspecified viral hepatitis, unspecified hepatitis, hepatitis.

[d] Sickle cell trait, thrombocytopenia, diseases of white blood cells, agranulocytosis, DiGeorges syndrome.

[e] Adverse effects of anti-infectives, tranquilizers, penicillin, and anti-rheumatics; erythema multiforme.

[f] Viral warts, acute gonococcal infection, unspecified urethritis, lymphogranuloma venereum, other venereal disease.

[g] Abscess of anal/rectal region, anal fissure or fistula, internal or external hemorrhoids.

[h] Hemorrhage of GI tract, idiopathic proctocolitis, unspecified disorder of rectum/anus, intestinal infection due to other organism (viral), unspecified gastritis, acute pancreatitis.

respectively. Among the infected men, 792 unique diagnoses were reported. Further analysis focused on only these diagnoses. Their frequencies ranged from 1 to 74 among infected men and from 0 to 189 among those uninfected. Of the 792 diagnoses, 120 (15%) had four or more occurrences among HIV-infected soldiers. Of these 120 conditions diagnosed four or more times, 79 (66%) were more frequent (PRs >1.0) among HIV-infected vs uninfected men. Of these, 21 diagnoses were nominally statistically more common among men later diagnosed with HIV infection and had PRs ≥2.0. Table 2 presents the PRs and 95% confidence intervals for these 21 diagnoses.

Table 3 presents a qualitative description of the major causes of morbidity among individuals in this study. Those diagnoses with elevated PRs in each of three time periods are shown. Two diagnoses, hepatitis B and abscess of the anal/rectal region, are more common over the entire 8-year follow-up period among individuals subsequently diagnosed with HIV infection. From 1 to 5 years prior to HIV diagnosis, conditions associated with increased risk of HIV infection (e.g., syphilis and sexually transmitted diseases [STDs] other than syphilis) and with early HIV disease (e.g., unspecified viral infection and enlargement of lymph nodes) are significantly more common. Finally, at 1 to 2 years prior to diagnosis, conditions possibly associated with acute retroviral syndrome (e.g., unexplained fever and mononucleosis) are seen more frequently.

Discussion

The AIDS epidemic was first recognized in the late 1970s and early 1980s when previously rare illnesses began to be reported among gay men in San Francisco and New York.[8,9] Since that time these illnesses have been found to be associated with late-stage HIV disease. The natural history of late-stage HIV disease has been well characterized.[10-12] However, there have been relatively few detailed characterizations of the early stages of HIV disease.

Current interest in early intervention in the HIV disease course has focused interest on its presymptomatic stages.[13,14] Recent work has shown the efficacy of antiviral and prophylactic therapeutic intervention.[15,16] Trials are underway to assess effects of therapies when given in early, asymptomatic stages of HIV disease. It is important to clearly understand the early natural history of this disease before therapies are made available for general use.

It is important to document the usual course of HIV disease for a number of reasons. First, the early course is often thought of as presymptomatic. Indeed, individuals in late stages of HIV disease are likely to have trouble recalling illnesses that occurred years earlier. Furthermore, when a diagnosis is made early in the clinical course, interventions are applied that are specifically meant to delay or prevent the presentation of signs and symptoms. Medical surveillance databases that can document morbidity and HIV-infection status are not commonly available. Populations that are routinely screened for HIV and that have access to medical care in a single system are uncommon. Finally, it is difficult to determine whether morbidity that occurred years prior to a diagnosis of HIV disease was associated with the pathophysiology of HIV disease or was part of the background of morbidity of the base population.

In this study we investigated the natural history of HIV disease using routinely and systematically collected hospital admission records from members of a defined base population. This method allowed investigation of the period prior to the time of each individual's HIV-infection diagnosis. Thus, the "true" (albeit partial) natural course could be described among individuals not yet aware of and, therefore, not yet receiving care specifically for their HIV infection. Since this method does not rely on recall, it is also possible to assess HIV-related morbidity, as well as other types of morbidity, including that related to behaviors associated with HIV-infection risk. Finally, this approach provides insight into morbidity in relation to the experience of uninfected individuals from the same base population.

Our results indicate that hospitalization rates significantly increased among men who were subsequently diagnosed with HIV infection. This experience was in contrast to that of a referent group. During each of the 8 years prior to the date of initial HIV-infection diagnosis, men who became HIV infected experienced significantly higher rates of inpatient morbidity than men who did not.

This increase may be explained by (1) morbidity associated with acute HIV infection (e.g., fever, mononucleosis), (2) morbidity associated with progression of HIV disease through its clinical course, and (3) morbidity associated with behaviors that increase risk of HIV infection (e.g., STDs).

TABLE 3—Diagnoses with Elevated Prevalence Ratios (PRs) over Three Time Periods Prior to HIV-Infection Diagnosis.[a]

PERIOD 1: 1 to 2 Years Prior to HIV-Infection Diagnosis	PERIOD 2: 3 to 5 Years Prior to HIV-Infection Diagnosis	PERIOD 3: 6 to 8 Years Prior to HIV-Infection Diagnosis
Hepatitis B	Hepatitis B	Hepatitis B
Abscess of anal/rectal region	Abscess of anal/rectal region	Abscess of anal/rectal region
Unspecified viral infection	Unspecified viral infection	Hepatitis A
Enlargement of lymph nodes	Enlargement of lymph nodes	
Syphilis	Syphilis	
Sexual transmitted diseases (not including syphilis)	Sexual transmitted diseases (not including syphilis)	
Chronic tonsillitis	Chronic tonsillitis	
Fever of unknown origin	Streptococcal sore throat	
Mononucleosis	Essential hypertension	
Unspecified pneumonia	Acute tonsillitis	
Acute pharyngitis	Anal fissure	
Deviated nasal septum	Lumbago	

[a] Inclusion criteria for any diagnosis into the three time periods were the occurrence of that diagnosis (1) at least three times in 1 year or (2) at least two times in 2 or more years per time period among individuals later diagnosed with HIV infection.

Since exact dates of infection with HIV are unknown, the experience of the HIV-infected group cannot be further subdivided into pre- and post-HIV-infection periods. However, since HIV infection is irreversible and antibody to HIV is persistent, an "exposure gradient" can be used to characterize their experience: at the RD, 100% were infected with HIV (by definition); at an unknown time in the past, 0% were infected with HIV; in between there is a monotonic gradient of the proportion of the group with HIV infection. While at a particular point the proportion with HIV infection is not precisely known, the experience of the HIV-Ab positive group near the RD is more associated with HIV infection than any experience more remote.

Although small sample sizes for most diagnoses did not allow for year-specific analyses, we were able to observe some trends over the 8-year period. For example, lifestyle-related morbidity (e.g., STDs, anal/rectal trauma) declined as the RD was approached. It is possible that individuals diagnosed with HIV infection became aware of their risk and changed behaviors and partner selection practices prior to diagnosis. Decreasing STD rates from the early to late 1980s have been reported,[17–20] particularly among White homosexual men. These decreasing trends have been variously ascribed to heightened awareness of and fear of contracting HIV/AIDS. In this study, even at their lowest point (1–2 years prior to HIV-infection diagnosis) lifestyle-related diagnoses occurred nearly four times more often in men subsequently diagnosed with HIV. Current studies assessing risk factors for HIV infection in recently diagnosed soldiers[21] may shed light on the relationships of STDs and other lifestyle-related morbidity on risk of HIV infection.

Overall, the amount of substance use–related morbidity was not different between HIV-Ab positive and HIV-Ab negative individuals. Alcohol abuse or dependency accounted for approximately 50% of substance use–related diagnoses in both groups. The remaining morbidity was accounted for by opioid, cannabis, cocaine, barbiturate, hallucinogen, and unspecified drug abuse. Notably, there were eight opioid-dependency diagnoses among men diagnosed with HIV infection compared to none among other men. This suggests that intravenous—but perhaps not other—drug use was a factor in these infections. Studies of behaviors among recent HIV seroconverters[21] may provide insights into trends of intravenous drug use in relation to HIV infection in military populations.

St. Louis and colleagues[22,23] recently estimated community HIV prevalence based on seroprevalences among patients admitted to hospitals for non-HIV–associated diagnoses. They concluded that hospital patients in certain geographic areas of the United States should be routinely screened for HIV infection. The results of our study may be useful in broader settings to guide the clinical judgment of care providers regarding the potential risk of HIV infection among patients. A review of patient medical history and personal practices may refine assessments of risk, encourage HIV diagnoses, and focus counseling and health education efforts in routine medical practices. As a result, HIV-directed prevention/intervention programs aimed at early stages of HIV infection could have a significantly greater impact on decreasing morbidity and health care needs than just those interventions aimed at treatment and care of end-stage HIV disease (e.g., ARC, AIDS). Thus, communities may have sufficient reason to implement comprehensive prevention programs, the impact of which might be felt prior to the time when AIDS is clinically and epidemiologically apparent.

It should not be surprising that prior history of STDs, acute viral illnesses of unknown etiology, and hepatitis B were more frequent among men subsequently diagnosed with HIV infection than others. We suggest that these diagnoses should alert primary care providers of the increased risk of HIV infection, particularly in settings such as STD, emergency, and methadone maintenance clinics. Patients with sentinel diagnoses should be screened for HIV infection, counseled regarding risks of infection and transmission, and, if negative, actively followed for reevaluation. ☐

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Utilization of Health Services in a Cohort of Intravenous Drug Users with Known HIV-1 Serostatus

ABSTRACT

Background. Intravenous drug users (IVDUs) with human immunodeficiency virus (HIV) infection and AIDS often have no health insurance or rely on public programs to finance their health care. We examined the independent contributions of HIV serostatus, clinical symptoms, CD4 cell counts, and health insurance to utilization of health services among 1881 intravenous drug users in Baltimore, Maryland.

Methods. Participants in an ongoing natural history study of HIV were informed of HIV serostatus and seropositives were informed of CD4 cell counts; 6 months later, participants were administered a questionnaire concerning self-reported use of health services, insurance coverage, and HIV-related symptoms.

Results. Of 1881 participants, 67% had health insurance (including Medicaid), 48% had at least one outpatient visit, and 12% had at least one inpatient visit within the prior 6 months. The proportion of the study population that was HIV-1 seropositive was 32%. In multivariate analysis, the single most important predictor of health care utilization was the presence of two or more HIV-related clinical symptoms. HIV positive serostatus alone or known low CD4 counts were not significantly associated with use of health care services.

Conclusions. These data suggest that HIV seropositive IVDUs are not receiving recommended preventive care. Additional efforts will be needed to ensure that HIV-seropositive drug users participate in currently recommended protocols for early treatment of asymptomatic HIV-1 infection. (*Am J Public Health.* 1991;81:1285–1290)

Liza Solomon, DrPH, Richard Frank, PhD, David Vlahov, PhD, Jacquie Astemborski, MHS

Introduction

By the end of January 1991, more than 161 000 cases of acquired immunodeficiency syndrome (AIDS) had been reported in the United States (K. Taylor, CDC AIDS monthly surveillance report, personal communication). Black and Hispanic adults account for a disproportionately large number of those infected.[1–5] The overrepresentation of minorities among those with HIV infection and AIDS is thought to be largely a consequence of intravenous (IV) drug use. It is estimated that IV drug use alone or IV drug use associated with homosexual or bisexual practices is responsible for 48% of the cumulative number of AIDS cases (K. Taylor, CDC, personal communication). One important outcome of the shifting demographics of the epidemic is that responsibility for financing and providing health care for people with human immunodeficiency virus type 1 (HIV-1) infection increasingly falls to the public sector.[6–9] Public policy will therefore play a significant role in influencing the availability of care, the timing of treatment, and the patterns of health service utilization. This research provides an initial examination of health services utilization among a population of IV drug users (IVDUs). Recent clinical research has demonstrated the important role of preventive care in delaying and mitigating the effects of opportunistic infections associated with HIV-1 disease.[10] Of concern is whether the recent advances in medical treatment, as well as the benefits of early clinical intervention, are available to the legions of IVDUs who comprise the second wave of the HIV epidemic.

The present study makes use of a data base available through an ongoing natural history study of HIV infection in IVDUs. The findings describe the relationship between health status and the utilization of medical services among IVDUs, as well as the role of health insurance in determining utilization patterns within this population.

Methods

Respondents to the health care utilization survey came from the AIDS Link to Intravenous Experience (ALIVE) Study, an ongoing longitudinal natural history study of HIV infection and IV drug use. ALIVE study participants were recruited through a number of strategies. Brochures describing the study were distributed to IVDUs who were being seen in agencies such as the Maryland Division of Parole and Probation, drug treatment programs, sexually transmitted disease clinics, emergency rooms, and homeless shelters. Also, participants were recruited via word-of-mouth. Eligibility criteria for the study included age of 18 years or older and a history of IV drug use in the previous 10 years. Individuals volunteering for the study were asked a series of questions concerning drug use history, incarcera-

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Liza Solomon, David Vlahov, and Jacquie Astemborski are with the Department of Epidemiology of The Johns Hopkins School of Hygiene and Public Health. Richard Frank is with the Department of Health Policy and Management of The Johns Hopkins School of Hygiene and Public Health.

Requests for reprints should be sent to Dr. Liza Solomon, Dept. of Epidemiology, The Johns Hopkins School of Hygiene and Public Health, 624 North Broadway, Room 763, Baltimore, MD 21205.

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TABLE 1—Sociodemographic and Drug-Use Characteristics of 1881 Intravenous Drug Users, by HIV-1 Serostatus, Baltimore, Md

Characteristic	n	% HIV-1 Positive
Age (years)		
18–24	154	29.2
25–34	896	36.0
35–44	689	30.6
≥45	142	21.8
Race[a]		
Black	1723	34.2
Non-Black	156	12.8
Gender		
Male	1491	31.5
Female	390	36.1
Legal income[a]		
<$5000	1380	33.1
$5000–$10 000	272	26.8
>$10 000	227	34.8
Education[b]		
1–8 grade	160	24.3
9–12 grade	1458	33.1
>12 grade	262	33.6
Homeless within last year		
Yes	644	32.3
No	1237	32.5
Employed at baseline[b]		
Yes	428	28.3
No	1452	33.7
Drug use		
Current injector	1411	32.6
No use in past 6 months	470	31.9
Insurance in past 6 months		
Medicaid	1085	60.5
Private	181	9.2
No insurance	615	30.3

[a] Missing 2.
[b] Missing 1.

tion, and history of homelessness in order to obscure the true eligibility requirements. Of the individuals who came to the study, 397 were disqualified because they did not admit to IV drug use. At entry into the study, 92% of the cohort demonstrated needle track marks on the extremity used for venipuncture. Participants reported a mean of 13 years of IV drug use at time of entry into the study. Forty-two percent of study participants reported a history of drug treatment within the previous 10 years and 13% reported a drug or alcohol treatment admission within the previous 6 months. The rationale and details of the data collection have been described in detail elsewhere.[11,12]

At the initial visit, after informed consent and pretest counseling procedures, trained interviewers administered a detailed questionnaire including sections on sociodemographic information, medical history, including symptoms associated with HIV disease (Appendix A), and current and past drug use and sexual practices.

All participants had blood drawn for HIV-1 antibodies. Specimens were determined to be positive if they were reactive on two ELISA tests (Genetic Systems, Seattle, Washington) and were confirmed with Western blot assay (DuPont, Wilmington, Delaware). All screening clinic participants were given AIDS risk reduction education and appointments to return to the clinic in 3 weeks to receive HIV test results. Participants were informed of the availability of a full-time addictions counselor on staff who would facilitate referrals for drug treatment. Study participants were able to see the addictions counselor on an immediate walk-in basis.

At the return visit, all seropositive individuals and a sample of 150 HIV-negative individuals were invited to participate in the follow-up clinic, which involved semiannual visits for behavioral interviews, physical exams, and venipuncture to assess CD4 counts.[13] All other study participants were invited to continue in the screening clinic, which consisted of semiannual HIV testing and interviews. At all points in the study, health was monitored, and HIV-positive individuals were given counseling and referred to the local HIV clinic for treatment or prophylaxis when indicated by symptoms or CD4 levels.

Beginning in December 1988, health care insurance and utilization questionnaires were administered to all individuals participating in the screening or follow-up components of the ALIVE study. The data presented below describe health care utilization, demographics, health status, and health insurance coverage of a group of 1881 study subjects. The validity of self-reports of insurance status was examined by comparing self-reported Medicaid coverage and Medicaid claim files on a subsample of the study population. An 88% agreement between self-reports and Medicaid files was found.

Statistical Analysis

Our analytic point of departure is a set of descriptions of the sample in terms of their use of health care services, HIV-1 serostatus, insurance coverage, and clinical symptoms. Two measures of severity of illness, presence of two or more AIDS-related symptoms, and CD4 levels were used. The selection of stratification levels for CD4 counts was chosen to conform to old and new guidelines for clinical services. Prior to 1990, a CD4 count <200 was used as an indication for initiation of prophylactic treatment with zidovudine,[9] whereas current practices call for the use of zidovudine when CD4 levels are <500.[14] Frequency distributions of variables by HIV-1 serostatus were compared using χ^2 statistics; continuous variables were compared by subgroups using analysis of variance.[15]

The independent impacts of service utilization on the three variables of primary interest were accomplished via estimation of a probit regression for two dichotomous dependent variables: whether or not any inpatient care was used by an individual during a 6-month period and whether or not any outpatient care was received during the same period. Obtaining unbiased estimates of the impact of insurance coverage on use of services is complicated because there may be unmeasured factors that determine both insurance coverage and service utilization, for example, the degree of concern with one's health. We used an instrumental variables estimation technique described by Maddala[16] to address this potential source of bias. The standard errors of the estimates for this two-stage instrumental variables model were calculated by following the procedure proposed by Amemiya.[17]

Results

Descriptive Findings

Table 1 presents sociodemographic features of participants in the health utilization study (N = 1881). This group was predominately male (79%), non-White (92%), of low socioeconomic status (income below $5000—73%), and had a median age of 34. The proportion of the study population that was HIV-1 seropositive was 32% (n = 610).

Table 2 shows the association between insurance in the 6 months prior to interview and HIV-1 antibody status. HIV serostatus was not significantly associated with the presence of any type of health insurance.

Table 3 shows comparisons for the number of inpatient and outpatient medical visits between HIV seropositive and seronegative individuals. HIV seropositive individuals averaged significantly more inpatient admissions (0.21) in 6 months compared to HIV seronegative individuals (0.13). No significant difference was observed in the mean number of outpatient medical visits or inpatient drug or alcohol admissions by HIV-1 serostatus.

The effect of severity of illness on health service utilization was examined via two separate measures of severity, CD4 levels and presence of two or more HIV-related symptoms. Table 4 shows comparisons between use of health services, CD4 counts, and HIV-related symptoms. The CD4 counts of HIV positive individuals were stratified into three groups—<200, 200–500, and >500—and the mean number of visits was compared to that observed in a group of HIV seronegative IVDUs. HIV seropositive individuals with CD4 counts <200 and those with counts of 200–500 were significantly more likely to use inpatient services than a comparison group of HIV negative drug users ($P < .021$ and P .033, respectively). HIV positive individuals with CD4 levels <200 had higher numbers of outpatient visits than the comparison group of HIV seronegative IVDUs ($P < .079$). There was no significant difference in number of outpatient visits noted between HIV seropositive individuals with CD4 counts >200 compared to number of visits of HIV seronegative individuals.

The effect of symptoms on utilization of health services is also presented in Table 4. The mean number of outpatient visits of HIV seronegative individuals was compared to both HIV seropositive individuals with zero or one symptom and HIV positive individuals with two or more AIDS-related symptoms. Those with two or more symptoms were found to have significantly more inpatient admissions and outpatient visits than HIV seronegative individuals ($P < .001$ and $P < .003$, respectively).

Table 5 shows the relationship between insurance coverage and utilization of medical services in the prior 6 months. All individuals with health insurance, regardless of their HIV serostatus, had significantly more inpatient admissions, outpatient visits, or drug or alcohol treatment admissions than those without health insurance. We also noted a significant relationship (not shown) between utilization of emergency rooms and absence of health insurance ($P < .001$).

TABLE 2—Proportion of Intravenous Drug Users with Insurance and Health Service Use in Prior 6 Months Stratified by HIV-1 Serostatus

	HIV-1 Serostatus	
Characteristic	Positive (n = 610)	Negative (n = 1271)
Insurance status in past 6 months		
Medicaid	60.5	56.3
Private or other insurance	9.1	9.8
None	30.3	33.8

Overall $P = .227$

TABLE 3—Mean (±SD) Number of Medical Visits in Past 6 Months Stratified by HIV-1 Serostatus, Baltimore, Md

	Serostatus		
Variable	Positive n = 610	Negative n = 1271	*P* Value
Inpatient admissions	0.21 (0.55)	0.13 (0.48)	0.003
Outpatient visits	1.90 (5.05)	1.61 (4.07)	0.216
Inpatient drug or alcohol treatment	1.7 (5.04)	1.4 (4.65)	0.23

TABLE 4—Analysis of Variance for Medical Service Utilization among Intravenous Drug Users by HIV-1 Serostatus, CD4 Count, and Symptoms, Baltimore, Md

Variable	n	Mean Number Visits over past 6 Months	*P* value
Inpatient admission			
HIV–[a]	1271	0.14	—
HIV+, CD4 >500[b]	275	0.16	0.364
HIV+, CD4 200–500	207	0.21	0.033
HIV+, CD4 <200	34	0.56	0.021
HIV–[a]	1271	0.14	—
HIV+, 0–1 symptoms[c]	459	0.15	0.454
HIV+, ≥2 symptoms	92	0.57	0.001
Outpatient visit			
HIV–[a]	1271	1.61	—
HIV+, CD4 >500[b]	275	1.94	0.359
HIV+, CD4 200–500	207	1.74	0.612
HIV+, CD4 <200	34	3.61	0.079
HIV–[a]	1271	1.61	—
HIV+, 0–1 symptoms[c]	459	1.73	0.654
HIV+, ≥2 symptoms	92	3.41	0.003

[a] Reference group.
[b] 94 missing CD4 cells.
[c] 59 missing symptoms.

Multivariate Findings

Table 6 presents estimates from the two-stage instrumental variables probit model. (Estimates from the first stage of the model are presented in Appendix B. The rationale for the instrumental model specification is presented in Appendix C.) We expected to find a positive correlation between HIV positive serostatus, older age, symptoms, and insurance with both inpatient and outpatient health service utilization. Only one measure of severity—presence of two or more AIDS-related symptoms—was entered into the model. This measure was chosen instead of CD4 counts because data were available for all study participants and CD4 counts of <200 were strongly cor-

TABLE 5—Type of Health Service Utilization in Past 6 Months Stratified by HIV-1 Serostatus and Insurance

	Visit		No Visit		
Health Service	n	% with insurance	n	% with insurance	P value
Inpatient admission					
HIV positive	98	83.7	512	67.0	<.001
HIV negative	137	83.9	1134	64.0	<.001
Outpatient visit					
HIV positive	310	80.7	300	58.3	<.000
HIV negative	604	77.3	667	56.1	<.000
Inpatient detox					
HIV positive	73	87.7	532	66.9	<.001
HIV negative[a]	168	85.1	1099	63.2	<.000

[a]Missing 4.

TABLE 6—Final Probit Regression Estimates of Inpatient and Outpatient Utilization[b]

Variable	Inpatient Use[a]	Outpatient Use[a]
HIV-1 seropositive	0.127 (1.44)	0.044 (0.651)
Age 1 (25–34)	0.300* (1.62)	−0.198* (1.64)
Age 2 (35–44)	0.274 (1.03)	−0.101 (0.55)
Age 3 (≥45)	−0.011 (0.25)	0.06[illegible] (0.27)
Male gender	−0.318 (0.99)	0.070 (0.29)
Black race	0.204 (1.31)	0.091 (0.76)
Education (9–12)	0.090 (0.63)	0.809 (0.72)
Education (>12)	−0.061 (0.36)	0.189 (1.41)
Income ($5 000–$10 000)	−0.871 (0.69)	−0.163* (1.81)
Income ($10 000)	−0.208 (1.43)	−0.191* (1.81)
Symptoms of AIDS (≥2)	0.687** (4.77)	0.268** (2.19)
Insurance[a]	0.649* (1.63)	0.322 (1.12)
Intercept	−1.869** (8.61)	−0.178 (1.18)
Loglikelihood	−643.36	−1262.3

Note. Reference groups: age <25, female, non-Black, < 9 years.
[a]Endogenous variable.
[b]Asymptotic t statistics in parentheses.
*p < .10.
**p < .05.

related with the presence of two or more HIV-related symptoms. The first column of the model presents the results for the determinants of inpatient utilization, and the second column presents the results for outpatient utilization as a dichotomous variable. Although the results of the model for inpatient and outpatient are qualitatively similar for both models of utilization, there are some important differences. In terms of the variables of central concern, the results indicate that holding constant demographic factors and AIDS-related clinical markers, having insurance was marginally associated with the probability of inpatient care ($P <$ .06). However, insurance coverage has no significant impact on the use of outpatient care.

The variable representing HIV status was estimated to have a positive though not significant association with the probability of using either type of health care services. Although the coefficient estimate in the inpatient utilization model had a t statistic of 1.44, indicating a moderate amount of precision in the estimate, this was not significant by conventional standards. This suggests that holding constant demographic factors, insurance coverage, and symptoms, HIV serostatus did not have a significant independent impact on service utilization. Formal tests for interactions between HIV positive serostatus and the presence of health insurance showed no significant effects. As Table 6 shows, after adjusting for demographic variables, insurance, and HIV serostatus, the presence of two or more AIDS-related symptoms demonstrated the strongest association with the likelihood that an individual would get either inpatient or outpatient care.

Discussion

This study examined the relationship between health insurance, HIV-1 serostatus, and AIDS-related symptoms and the impact of these factors on the utilization of both outpatient and inpatient medical, drug, and alcohol treatment services. The presence of two or more AIDS-related symptoms had the strongest influence on the use of both inpatient and outpatient health care services. This is consistent with other research documenting the important role of symptoms and complaints in predicting service use.[18,19] In the multivariate analysis neither HIV-1 positive serostatus nor low CD4 levels in the absence of symptoms were associated with increased health services utilization. However, health insurance was significantly associated with the likelihood of using inpatient medical, drug, or alcohol treatment services. Even among individuals with low CD4 counts or AIDS-related symptoms, the absence of health insurance was associated with lower probabilities of receiving inpatient treatment. We did not find a significant relationship between the presence of health insurance and outpatient service utilization, which may be due in part to the availability of free care in East Baltimore. For example, the HIV clinics at certain local hospitals treat all individuals regardless of insurance status.

The next major finding is a negative one. Asymptomatic HIV seropositive subjects were estimated to be no more likely to use services than are HIV seronegative individuals. This is a particularly important problem because the greatest advances in treating HIV disease involve the postponement of manifestation of opportunistic infections that are the major causes of death in AIDS patients. Current medical practice recommends prophylactic treatment when CD4 counts drop below 500 mm^3, a level at which symptoms are not usually present. Failure to enroll HIV-positive individuals in health care prior to the appearance of symptoms suggests that preventive care, monitoring of CD4 counts, zidovudine prophylaxis, and prophylaxis for pneumocystis pneumonia are probably lacking and implies that individuals are not seeking care until they become very ill.

Programs to expand testing for HIV-1 in high-risk groups have been endorsed as an important element in promoting the early treatment of HIV-1 disease.[20] However, this study suggests that identification of HIV-1 positive status alone does not necessarily result in the initiation of preventive care. IVDUs often remain outside the health care system, using services on an emergent basis and frequenting emergency rooms rather than primary care clinics. Characteristics inherent in drug addiction (e.g., an overwhelming need to secure drugs daily) may make IVDUs particularly unable to conform to the requirements of many clinics for scheduled appointments and compliance with therapeutic plans. Efforts to encourage the early use of preventive services in this population will need to confront both the preferences and habits of the IV-drug-using population as well as institutional impediments to the use of services.[21,22]

Additionally, attention must be focused on making the drug-using community aware of the importance of early and continuous medical assessment and care. To date, education concerning HIV has focused on prevention of infection with implicit messages concerning the fatal consequences of HIV disease. Existing programs may have overlooked the need to provide information about the benefits of preventive treatment. Programs to expand medical services to drug treatment programs have become more common in Baltimore. Although this may increase knowledge and use of medical services to individuals participating in these drug treatment programs, our research suggests that the vast majority of drug users remain outside the drug treatment setting.

Conclusions based on this study may be limited by characteristics of the study population. This study reimbursed individuals for their participation; therefore, those with acute financial need may be overrepresented. Additionally, reports of insurance status as well as health services utilization were based on self-reports, the accuracy of which may be questionable. In spite of these potential limitations this study presents data on a cohort of drug users that are largely unaffiliated with a drug treatment program. Thus it may offer a more accurate picture of the use of services and the extent of care that many HIV seropositive drug users outside of treatment facilities are receiving. This study suggests that in spite of recommendations from the Centers for Disease Control for early treatment of HIV disease, many drug users still remain outside the health care system without benefit of preventive care. ☐

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APPENDIX A—HIV-Related Symptoms

Tired all the time for 2 or more weeks
Shortness of breath for 2 or more weeks
Diarrhea for 2 or more weeks
Unintentional weight loss of 10 or more pounds
White patches in the mouth for 2 or more weeks

APPENDIX B—First Stage Probit Regression Estimates for Insurance

Variable	Insurance[a]
HIV positive	0.73 (1.05)
Age 1 (25–34)	0.154 (1.32)
Age 2 (35–44)	0.495* (4.07)
Age 3 (≥45)	0.914* (5.35)
Male	0.722* (7.99)
Black	−0.162 (1.41)
Education 1 (9–12)	−0.058 (0.49)
Education 2 (>12)	−0.032 (0.22)
Income 1 ($5,000–$10,000)	−0.061* (0.69)
Income 2 (>$10,000)	0.226 (2.20)
Children	0.097 (1.41)
Employed	−0.240* (3.13)
AIDS symptoms	0.245* (2.45)
Intercept	0.041 (0.25)
Loglikelihood	−1067.3

Asymptotic *t* statistics in parentheses.
*p ≤ .05.

APPENDIX C—Rationale for Instrumental Variables Model Specification

Within the model specification it is important to include at least some independent variables that are not included as part of the utilization model. While this is difficult to accomplish on a priori grounds, we choose two variables that relate to the price of insurance. This is consistent with sequential budgeting on the part of consumers. Under such arrangements consumers first make a choice regarding the purchase of insurance and then subsequently choose to buy other goods based on the remaining income. Under such a model, variables affecting the price of insurance will have little effect on utilization. The two variables that differ between the two models are whether or not individuals have children and whether or not they are employed. We posit that these variables have a strong impact on the likelihood of having insurance and little effect on utilization once income, demographic features, and health status are taken into account.

Risk Factors for Shooting Gallery Use and Cessation among Intravenous Drug Users

David D. Celentano, ScD, David Vlahov, PhD, Sylvia Cohn, MPH, James C. Anthony, PhD, Liza Solomon, DrPH, and Kenrad E. Nelson, MD

ABSTRACT

Background. Shooting galleries, locations where intravenous drug users (IVDUs) can rent or borrow needles and syringes, are a high-risk environment for HIV-1 transmission. This study investigates risk factors for lifetime attendance at shooting galleries and differentiates characteristics of those who continue to frequent shooting galleries and those who have stopped.

Methods. We interviewed 2615 active IVDUs in Baltimore in 1988 and 1989 and determined patterns of IV drug use, sociodemographics, and HIV-1 serostatus as related to persistence vs cessation of shooting gallery use.

Results. Over half (52%) of active IVDUs reported ever using a shooting gallery, with 33% reporting use within the prior 3 months. In multivariate analysis, lifetime shooting gallery use was associated with male gender, homosexuality/bisexuality, low socioeconomic status, Black race, and heavier drug involvement. Persistent shooting gallery users were more frequently male, homosexual/bisexual, homeless, less educated, and started IV drug use more recently compared with those who ceased going to shooting galleries.

Conclusions. Shooting gallery attendance may be pragmatic from a sociological and economic perspective, but it carries with it a heightened risk of acquiring HIV-1 infection. (*Am J Public Health*. 1991;81:1291–1295)

Introduction

Behaviors that increase the risk of infection with human immunodeficiency virus, type 1 (HIV-1) among intravenous drug users (IVDUs) are primarily those involving transfer of blood through shared injection equipment.[1-5] Use of shared equipment at shooting galleries is particularly high risk for acquiring HIV-1 infection.[6-8] Shooting galleries are locations, generally clandestine, where IVDUs can rent, borrow, or purchase needles and syringes; frequently the injection equipment is used repeatedly and may contain contaminated blood. Repeated rental of syringes results in a sequential anonymous sharing of potentially contaminated equipment. Such sharing facilitates the transmission of blood-borne pathogens.[1]

Although needle sharing has been empirically demonstrated to be one of the principal AIDS risk behaviors among IVDUs,[9,10] reports on the characteristics and behaviors of IVDUs who use shooting galleries are sparse. Ethnographic reports suggest that IVDUs go to shooting galleries for several reasons, including (1) the illegality of possession of drug injection equipment and consequent fears of arrest; (2) the scarcity of syringes, leading to an economic imperative to share injection equipment; and (3) the social context of intravenous drug use.[11,12] In this report, we used social survey methods on a large sample of currently active IVDUs to complement the findings from the published ethnographic reports to examine reasons for shooting gallery use and to identify reasons for cessation vs persistence in gallery use.

Methods

Sample

Between February 1988 and March 1989, IVDUs in Baltimore, Md, were recruited into a longitudinal investigation of the natural history of HIV-1 infection [the AIDS Links to the Intravenous Experience (ALIVE) Study]. Recruitment relied predominately on word-of-mouth by community outreach efforts (85%); referrals from drug abuse treatment programs, sexually transmitted disease clinics, local emergency departments, university hospital HIV-1/AIDS clinics, and shelters for the homeless also assisted with cohort recruitment. In order to be eligible for enrollment, IVDUs had to be age 18 or older and have self-reported illicit drug injection in the prior 10 years. The rationale and details of the sampling and data collection plan have been described elsewhere.[13]

A total of 2921 participants were enrolled in the ALIVE study and completed baseline interviews and venipuncture. In this analysis, we restricted the sample to the 2615 IVDUs who reported injecting illicit drugs within 6 months prior to interview.

David D. Celentano is with the Department of Health Policy and Management, David Vlahov, Sylvia Cohn, Liza Solomon, and Kenrad E. Nelson are with the Department of Epidemiology, and James C. Anthony is with the Department of Mental Hygiene, The Johns Hopkins School of Hygiene and Public Health, Baltimore.

Requests for reprints should be sent to David D. Celentano, Department of Health Policy and Management, The Johns Hopkins University, School of Hygiene and Public Health, 624 N Broadway, Room 750, Baltimore, MD 21205.

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Data Collection Procedures

After providing voluntary informed consent, participating IVDUs were assigned a unique and confidential identifier and agreed to venipuncture for serum for HIV-1 antibody assays. Next, participants were questioned by a trained interviewer in a private room concerning demographic information, medical history, drug injection practices since 1977 on an annual basis, 10-year history of sexual activity, knowledge and attitudes about HIV-1 infection and AIDS, and use of medical care. A detailed description of the data-gathering procedures appears elsewhere[14]; a brief overview is presented here.

The self-report demographic data analyzed here include age, gender and sexual orientation, race/ethnicity, education, current employment, receipt of public assistance, and occurrence of homelessness within the prior 10 years. Participants were asked to recall their drug use behavior in the 6 months prior to their last injection and also in the first 3 months of use. Drug injection frequency reported for the prior year was categorized as once or more a day or as less frequently. Other drug involvement factors analyzed included self-reports of treatment for drug abuse, being arrested and incarcerated within the prior 10 years, and duration of drug use as indicated by the year when drugs were first injected. Although there was a battery of questions on history of sexual practices, we restricted the analysis here to one indicator: whether any sexual partner was an IVDU.

We inquired as to the use of shooting galleries on an annual basis (thereby determining never and ever use) beginning in 1977 or year of first drug use. Recent shooting gallery attendance was ascertained as follows: "During the 6 months before the last time you injected drugs, how many times did you go to a shooting gallery and use works borrowed, rented, or bought there?" For the analyses here we aggregated the responses into three categories: never went to a shooting gallery, former shooting gallery user but no use in the prior 6 months, and current user of shooting galleries (i.e., attendance within the prior 6 months).

Analysis

To examine the associations of sociodemographic and drug use characteristics with attendance at shooting galleries, we cross-classified current, former, and never users of shooting galleries with the predictor characteristics, using χ^2 statistics as an aid to interpretation. The strength of associations between these characteristics and practices and shooting gallery attendance was estimated by the odds ratio (OR), using the group of IVDUs who reported never going to a shooting gallery as the reference category. Logistic regression was used to examine the simultaneous association of these factors to lifetime shooting gallery use (comparing ever vs never), as well as to identify factors associated with persistence of this risk behavior (comparing current and former users). We used 95% confidence intervals (CIs) as an aid to statistical interpretation of results.

Results

Of the 2615 IVDUs still injecting drugs enrolled in the ALIVE study, 32.9% (n=861) reported attendance at a shooting gallery within the prior 6 months, and 19.4% (n=507) reported having gone to a shooting gallery in the past but not within the prior 6 months. Slightly less than one half (47.7%; n=1247) of the respondents denied ever having injected drugs in a shooting gallery.

The relationships between the demographic and drug use characteristics and lifetime shooting gallery attendance are presented in Table 1. Those IVDUs who have ever used shooting galleries were more likely to be male, homosexual/bisexual, slightly older, Black, and unemployed, to have received public assistance, and to report being homeless sometime in the prior decade, as compared to IVDUs who reported never injecting drugs in a shooting gallery. A review of drug involvement factors shows shooting gallery attenders were more likely to report an arrest or incarceration history, to have been in drug treatment as compared to never users, and to have initiated injection behavior earlier in their lives. With respect to current drug use, shooting gallery users were more likely to be daily drug users and to report sexual partners who were also IVDUs. Table 2 compares the characteristics of current shooting gallery users and IVDUs who reported stopping shooting gallery use. Current shooting gallery attenders were, again, male, homosexual/bisexual, younger, less educated, more often homeless in the prior decade, and had a history of injection behavior showing more recent initiation into this culture. No other drug involvement factor differentiated those persisting in shooting gallery use and those who ceased this activity.

Table 3 presents the results of two multiple logistic regression analyses: (1) ever vs never use of shooting galleries and (2) current vs former shooting gallery users. Age was deleted so that year of first IV drug use might be included in the equations. Using multiple regression to characterize those IVDUs who reported ever going to a shooting gallery compared with those with no self-report history of gallery attendance, we found that ever attenders were more likely to be male, homosexual/bisexual, Black, and homeless, to report arrests, drug treatment, and sex partners who were IVDUs, and to have a history of daily drug injection and of starting drug injection earlier, all other variables held constant. Compared to females, males, especially homosexuals or bisexuals, more commonly visited shooting galleries. The comparison of characteristics differentiating individuals who persist in injecting illicit drugs in shooting galleries and those who have stopped showed fewer distinguishing factors once other variables were held constant using multiple regression: recent gallery attenders had somewhat less education, more frequently reported homelessness, and more recently initiated IV drug use. The high OR for male homosexuals or bisexuals who persist in using shooting galleries is of concern because they may serve as a bridge for HIV-1 transmission from gay males who do not inject to heterosexual IVDUs.

Discussion

These data indicate that shooting gallery use is associated with lower socioeconomic status, greater criminal involvement, and indicators of heavier drug use—characteristics congruent with prior ethnographic reports of street drug users.[15–17] Our findings suggest that the use of shooting galleries is in part a pragmatic response to the contemporary context of illicit drug use, namely, that IVDUs seek access to needles for injecting drugs that are illegal to possess.[18] Shooting galleries provide needles in relatively clandestine locations, also reducing the threat of arrest for drug use and drug and drug paraphernalia possession.[19,20] However, the lack of financial resources associated with shooting gallery use fails to differentiate IVDUs who have stopped using shooting galleries from those that persist in this behavior, with the exception of greater reports of homelessness among re-

TABLE 1—Percentage Distributions and Odds Ratio (OR) Estimates of Sociodemographic and Drug Use Practices on Lifetime Shooting Gallery Attendance

	Shooting Gallery Attendance			
	Never (n = 1247)	Ever (n = 1368)	OR	95% CI[a]
Gender/Orientation				
Female	23.7	14.1	1.0	
Male/Heterosexual	70.0	76.4	1.8	1.5–2.2
Male/Bisexual	6.3	9.5	2.5	1.8–3.5
Age (years)				
18–24	8.6	5.7	1.0	
25–29	22.9	16.6	1.1	0.8–1.6
30–34	29.0	29.6	1.5	1.1–2.2
35–39	21.1	26.6	1.9	1.3–2.7
≥40	18.4	21.5	1.5	1.2–2.5
Race				
White	12.9	6.7	1.0	
Black	87.1	93.3	2.0	1.6–2.7
Education (years)				
≤12	55.8	56.8	1.0	
>12	44.2	43.2	1.0	0.9–1.2
Currently employed				
Yes	23.7	19.2	1.0	
No	76.3	80.8	1.3	1.1–1.6
Public assistance				
No	35.0	25.7	1.0	
Yes	65.0	74.3	1.6	1.3–1.8
Homeless[b]				
No	57.3	45.6	1.0	
Yes	42.7	54.4	1.6	1.4–1.9
Arrested[b]				
No	30.2	17.0	1.0	
Yes	69.8	83.0	2.1	1.8–2.6
Incarcerated[b]				
No	40.3	27.3	1.0	
Yes	59.7	72.7	1.8	1.5–2.1
Drug treatment (ever)				
No	59.1	37.0	1.0	
Yes	40.9	63.0	2.4	2.1–2.8
Year initiated IV drug use				
1988	7.3	1.5	1.0	
1986–1987	13.0	6.0	2.3	1.3–4.0
1984–1985	8.5	5.0	3.0	1.7–5.2
1980–1983	15.9	12.1	3.8	2.3–6.5
Before 1980	55.3	75.4	6.8	4.2–11.1
Frequency of drug injection				
< Daily	69.2	51.3	1.0	
Daily	30.8	48.7	2.1	1.8–2.5
IVDU[c] sex partner				
No	35.1	17.5	1.0	
Yes	64.9	82.5	2.6	2.1–3.1

[a] CI = confidence interval.
[b] During the previous 10 years.
[c] IVDU = intravenous drug user.

cent shooting gallery users. Homeless individuals are more likely to frequent shooting galleries and to persist in this behavior, thereby providing some limited evidence supporting an hypothesis of economic necessity. The traditional view that needle sharing reflects a culturally entrenched ritual resistant to change[21] has been expanded to acknowledge a pragmatic response of IVDUs to the restricted availability of needles and syringes.[22] Lessons learned from studies of needle sharing offer some insights into the complexity of this social behavior but appear insufficient to fully explain its continued existence.

While going to shooting galleries may be pragmatic from a sociolegal and economic perspective, using needles obtained from shooting galleries is associated with increased risk of acquiring HIV-1 infection. The extent of exposure to HIV-1 in shooting galleries was unknown until recently. Chitwood et al.[23] obtained 148 needles and syringes from three shooting galleries in Miami, of which 10.1% tested positive for HIV-1 antibody. This relatively high rate of exposure found in shooting galleries may reflect frequent reuse of equipment as well as sharing equipment more frequently with strangers or acquaintances whose HIV-1 antibody status may be unknown (as compared to sharing needles among groups of close friends).[24,25] In our study, the odds of having antibody to HIV-1 among those who have ever gone to a shooting gallery were nearly twice the odds for those who had reported never attending a gallery.[8] Vigorous prosecution of drug paraphernalia offenders may act to facilitate transmission of HIV-1 infection rather than reduce its spread.

One public health approach to the problem of HIV-1 transmission in shooting galleries is to offer sterile needles through needle exchange programs.[26-28] This approach generally has been considered incompatible with current policies on drug use in most US jurisdictions. A recent report[7] discusses public health preventive interventions to reduce the risk of HIV-1 among users of shooting galleries in the context of the current legal climate—educating proprietors of shooting galleries to educate their clients and to improve needle sterilization practices. This approach requires the active participation of shooting gallery operators in creating "safer" galleries; additional studies will be required to determine the extent to which such interventions are feasible and effective. Ultimately, a goal for reducing the transmission of HIV-1 infection among IVDUs would involve the cessation of shooting gallery use, rather than attempts to create "safer" galleries. Study results presented here suggest that although a large proportion (37%) of IVDUs reporting ever going to shooting galleries have recently stopped, the majority persist in this activity. Data on AIDS knowledge and attitudes posed to these participants demonstrated no trends differentiating these three groups of gallery users.

A focus on risk reduction among homosexual or bisexual males who inject drugs in shooting galleries seems to be of paramount importance. Although patterns of shooting gallery use fluctuate over time and our classification into three distinct groups based on recent history may be arbitrary, no distinct trends in use were detected as related to the risk factors assessed. Although other studies report behavior change among IVDUs in response to the HIV-1 epidemic,[29] these

TABLE 2—Percentage Distributions and Odds Ratio (OR) Estimates of Sociodemographic and Drug Use Practices on Current vs Former Shooting Gallery Attendance

	Shooting Gallery Attendance			
	Current (n = 861)	Former (n = 507)	OR	95% CI[a]
Gender/orientation				
Female	12.1	17.7	1.0	
Male/Heterosexual	76.0	77.2	1.4	1.1–2.0
Male/Bisexual	11.9	5.1	3.4	2.0–5.7
Age, y				
18–24	7.3	3.0	1.0	
25–29	17.4	15.2	0.5	0.3–0.9
30–34	29.3	30.2	0.4	0.2–0.7
35–39	24.3	30.6	0.3	0.2–0.6
≥40	21.7	21.1	0.4	0.2–0.8
Race				
White	6.4	7.3	1.0	
Black	93.6	92.7	1.1	0.8–1.6
Education, y				
≤12	59.4	52.5	1.0	
>12	40.6	47.5	1.3	1.1–1.7
Currently employed				
Yes	18.9	19.5	1.0	
No	81.1	80.5	1.0	0.8–1.4
Public assistance				
No	26.0	25.3	1.0	
Yes	74.0	74.7	1.0	0.8–1.2
Homeless[b]				
No	41.7	52.3	1.0	
Yes	58.3	47.7	1.5	1.2–1.9
Arrested[b]				
No	17.1	16.8	1.0	
Yes	82.9	83.2	1.0	0.7–1.3
Incarcerated[b]				
No	25.9	29.6	1.0	
Yes	74.1	70.4	1.2	0.9–1.5
Drug treatment (ever)				
No	36.5	37.9	1.0	
Yes	63.5	62.1	0.9	0.7–1.2
Year initiated IV drug use				
1988	2.2	0.2	1.0	
1986–1987	7.9	2.8	0.3	0.1–2.1
1984–1985	5.0	5.1	0.1	0.1–0.7
1980–1983	11.4	13.4	0.1	0.1–0.6
Before 1980	73.5	78.5	0.1	0.1–0.6
Frequency of drug injection				
< Daily	50.3	53.1	1.0	
Daily	49.7	46.9	1.1	0.9–1.4
IVDU[c] sex partner				
No	16.0	19.9	1.0	
Yes	84.0	80.1	1.3	0.9–1.7

[a] CI = confidence interval.
[b] During previous 10 years.
[c] IVDU = intravenous drug user.

studies note that a substantial proportion of IVDUs persist in risk-increasing behaviors.[30] Further, studies have shown that although some IVDUs who engage in needle sharing practices are knowledgeable about modes of transmission,[31,32] they do not perceive themselves to be at high risk for infection.[33] These data are consistent with previous reports[29,30] indicating that a large proportion of IVDUs continue to engage in high risk behaviors, despite their knowledge about HIV-1.

The characteristics of those who continue to use shooting galleries suggest that we need to target IVDUs earlier in their careers, before this behavior becomes firmly entrenched, and to focus additional interventions on an especially difficult-to-access subset of IVDUs—those who are less educated and impoverished. Paradoxically, those who persist in using shooting galleries are also likely to report a history of drug treatment, a characteristic shown to lead to reduced IV drug use[34] although relapse is frequent.[35] However, those who continue going to shooting galleries are not necessarily more frequent drug users than those who have stopped, suggesting that overall drug involvement is not a discriminating characteristic. HIV-1 prevention efforts must strive for greater access to IVDUs who use shooting galleries in order to appropriately target messages on cessation of shooting gallery use. Prevention efforts should also consider the approach suggested by Chitwood et al.[23] and articulated by Des Jarlais and Friedman[7] of making galleries safer for those who use them. ☐

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TABLE 3—Multiple Logistic Regression Analysis Results on Lifetime and Recent Shooting Gallery Attendance

	Ever vs Never		Current vs Former	
	OR	95% CI	OR	95% CI
Gender:				
Male heterosexual	2.1	1.7–2.7	1.8	1.3–2.5
Male homosexual/bisexual	2.5	1.7–3.6	4.1	2.4–7.2
Race: Black	2.6	1.9–3.5	1.3	0.8–2.1
Education: ≤12 years	1.1	0.9–1.3	1.3	1.1–1.7
Currently employed: No	1.0	0.9–1.3	1.0	0.7–1.4
Public assistance: Yes	1.2	1.0–1.5	0.9	0.7–1.3
Homelessness: Yes	1.4	1.1–1.7	1.6	1.2–2.0
Arrested: Yes	1.4	1.1–1.8	0.7	0.5–1.1
Incarcerated: Yes	1.2	0.9–1.4	1.2	0.9–1.6
Drug treatment: Yes	1.9	1.6–2.3	1.3	1.0–1.6
Daily injection: Yes	1.8	1.5–2.2	1.3	1.0–1.6
Year started IV drug use	1.3	1.2–1.4	0.7	0.6–0.8
Sex partner IVDU: Yes	2.3	1.9–2.8	1.3	0.9–1.7
Model χ^2	463.4, $P<.0001$		77.3, $P<.0001$	

Note. OR=odds ratio; CI=confidence interval; IVDU=intravenous drug user.

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Comparing the Predictive Accuracy of Health Risk Appraisal: The Centers for Disease Control Versus Carter Center Program

Julie A. Gazmararian, MPH, Betsy Foxman, PhD, Louis Tze-Ching Yen, PhD, Hal Morgenstern, PhD, and Dee W. Edington, PhD

ABSTRACT

Background. From 1986 to 1987 the Carter Center of Emory University joined with the Centers for Disease Control (CDC) to develop a new, probability-based, adult health risk appraisal (HRA) instrument for the public domain. This new instrument is compared with the CDC HRA version to determine which is a more accurate predictor of mortality.

Methods. We compared predicted mortality risks from the CDC HRA and the Carter Center HRA with the observed mortality among 3135 smokers and never-smoking persons, aged 25 to 60, followed from 1959 to 1979 as part of the Tecumseh Community Health Study.

Results. When individuals were classified according to the difference between their actual age and risk age, for the CDC HRA, there was a progressively increasing risk of 10-year mortality as the difference increased. The Carter Center HRA did not show this trend. An analysis using relative operating characteristic curves showed that the mortality risk predictions for both programs were very similar for men and women. However, differences between actual age and risk age for the two programs were not similar for men or women, particularly older men. Therefore, actual age minus risk age for the CDC program was a more accurate predictor of 10-year mortality than was this difference for the Carter Center program.

Conclusions. The results from both types of analyses suggest that the validity of risk ages obtained from the Carter Center version may not be sufficient to justify updating programs for those currently using the CDC instrument. (*Am J Public Health.* 1991;81:1296–1301)

Introduction

The health risk appraisal (HRA) has become a popular approach to help people identify the risks associated with personal characteristics (biological, life-style, family history).[1] Risk factors identified from various epidemiologic studies are combined with mortality statistics in order to assess an individual's risk of mortality within the next 10 years. An individual's risk is usually expressed as risk age (or appraised age), which enables one to compare one's own modifiable risk with that of a cohort. Thus the HRA could be a reasonably efficient method for transmitting this mortality risk information to individuals and stimulating them to change in terms of attitudes, beliefs, and life-styles, if so indicated.[2,3]

Although there are numerous variations of the HRA instrument, the majority are adapted from techniques developed by Robbins and Hall.[4] The accuracy of the mortality predictions from this methodology has been questioned repeatedly.[1–3,5–9] For example, Smith et al.[10] report that this method systematically overestimates the probability of mortality from coronary heart disease when the predictions are correlated with those from the Framingham Heart Study and the Risk Factor Update Project. Foxman and Edington assessed the accuracy of the Centers for Disease Control (CDC) HRA by comparing HRA-predicted risks of mortality with deaths that actually occurred among 3135 persons followed from 1959 to 1979 as part of the Tecumseh Community Health Study.[7] They found that the differences between actual age and CDC HRA risk age correctly classified individuals into high- and low-risk groups.

In an effort to update the scientific base of the HRA instrument, the Carter Center of Emory University joined with the CDC from 1986 to 1987 to develop a new, probability-based, adult health risk appraisal instrument for the public domain. The result of this project is "Healthier People," the Carter Center of Emory University Health Risk Appraisal Program.[11] The Carter Center HRA incorporates mortality tables from 1980 to 1982; integrates recommendations from the Breslow Risk Factor Update Project; uses regression equations from heart attack and stroke estimates developed by the Framingham Study; employs new models for cancer estimates developed from National Cancer Institute data; corrects height-weight recommendations; includes high-density lipoprotein as a risk factor; eliminates race as a predictor variable, so that mortality estimates are based only on age and sex; replaces the credit-debit method with multivariate statistical techniques; adjusts reference mortality projections for causes of death where risks are quantified by nonmodifiable precursors, so that individuals are not penalized if they have high levels of fixed risk, e.g., family history of breast cancer, diabetes, etc.; and calculates an individual's risk from 19, rather than 11, quantifiable causes of

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Julie A. Gazmararian and Betsy Foxman are with the School of Public Health at the University of Michigan. Dee W. Edington and Louis Tze-Ching Yen are with the Division of Kinesiology at the University of Michigan. Hal Morgenstern is with the School of Public Health at the University of California, Los Angeles.

Requests for reprints should be sent to Dr. Dee W. Edington, Division of Kinesiology, 3040C CCRB, University of Michigan, Ann Arbor, MI 48109-2214. This paper was submitted to the journal February 7, 1990, and accepted with revisions February 19, 1991.

death. The Carter Center promises continuous updating of software as new scientific findings are determined and evaluated and technical enhancements developed.[11,12]

We have compared this new HRA instrument (version 3.1, released in late 1988) with the CDC HRA version that was used in our previous study[7] to see if this new version more accurately predicts mortality. This comparison has important implications for those using the HRA instrument for scientific and research purposes as well as for those using it for health education or intervention.

Methods

As in our previous study,[7] data available from the Tecumseh Community Health Study were used as input to the HRA programs. In order to compare the two instruments, we calculated the average HRA-predicted 10-year mortality risk (predicted risks) and risk age for both programs.

Risk age is defined for a given individual as the expected age of someone in the reference population of the same age, sex, and other fixed characteristics for whom the average 10-year mortality risk is the same as the HRA-predicted 10-year mortality risk. Thus, for the CDC HRA method a risk age of 50 for a 40-year-old White male means that the 10-year mortality risk predicted for that person is equal to the life-table estimate of the 10-year mortality risk for a 50-year-old White male in the 1975 to 1977 US population (the reference population for the CDC HRA). Hence this 40-year-old person is in poorer health than someone of the same age, sex, and race in the reference population.

The Carter Center HRA compares the predicted risk to the cohort average risk mortality, i.e., the population average risk for a person of the same age and sex adjusted for the presence of nonmodifiable risk factors.[11] Thus an individual with a family history of breast cancer would be compared with a reference population different from that of an individual without a family history of breast cancer. This reference mortality adjustment is made only for causes of death where risks are quantified by nonmodifiable precursors (US population projections are used as the reference mortality for all other causes of death).[12] A detailed comparison between the CDC HRA and the Carter Center HRA calculation of an individual's predicted risk of mortality and risk age is outlined in Appendix A.

Data Instruments

The Carter Center HRA (version 3.1) computes health risks based on a 45-item questionnaire[11]; the CDC HRA uses a 37-item questionnaire.[13] Twenty-nine of the 45 questions for the Carter Center HRA and 31 of the 37 questions for the CDC HRA are used directly for computation of risk (predicted risks and risk age, Appendix B). These questions were selected by the program's developers for their presumed causal relationship rather than mere statistical association. When items are not answered, both programs assign population norms.

Study Population

The Tecumseh Community Health Study began in 1959 as a comprehensive prospective investigation of health and disease in a community.[14] The principal aim was to identify causes and precursors of coronary heart disease, hypertension, chronic respiratory disease, diabetes mellitus, and other chronic diseases. Mortality ascertainments were completed successfully on more than 99% of previously examined subjects on three occasions, the most recent being 1978 to 1979.

In the 20-year period following the first cycle of examinations, 1059 (12.3%) of the 8641 participants died. Since the HRA is considered most accurate in the 25- to 60-year age range, we limited our analyses to the 3762 persons aged 25 to 60 in 1959 to 1960. Time since stopped smoking was not available for the 564 former smokers but is required by the CDC HRA; excluding former smokers from the current study restricted our sample to 3198; vital status after 10 years was available for 3166, and 3135 had sufficient information to calculate HRA. Thus a total of 3135 never-smoking or currently smoking individuals aged 25 to 60 in 1959 were included in this analysis.

Of the 29 variables used by the Carter Center HRA in predicting risk age, 13 were collected during the 1959 to 1960 Tecumseh survey, and 14 of the 31 variables used by the CDC HRA were available (Appendix B). Some recoding of the Tecumseh survey responses (mainly collapsing of categories to correspond to the precoded responses of both HRAs) was necessary for six of the variables. Although the wording was somewhat different in the two HRA instruments, the meanings were judged to be the same.

Data Analysis

Using both programs for each individual from data gathered during the 1959 to 1960 Tecumseh Community Survey, we compared the predictions of the CDC HRA with the Carter Center HRA, calculating the average HRA-predicted 10-year mortality risk from all causes (predicted risks) and risk age. The difference between each individual's age at the baseline survey and risk age (age difference) was compared by gender and age groups. To determine if risk age was a good predictor of observed mortality, we examined the proportion dying at 10 years for both men and women by the difference between actual age and risk age. The average HRA-predicted mortality risk, observed mortality, and age-adjusted mortality risk were calculated for each category of this difference between actual age and risk age. Since variations in age, sex, and race explain much of the variability in death rates, examining the gender, age, and race-specific differences between actual age and risk age represents the contribution of nondemographic inputs to the HRA.[15]

In order to assess which instrument more accurately predicted mortality, we performed relative operating characteristic (ROC) curve analysis[16] to compare both the predicted risks and the difference between actual age and risk age for each prediction method. ROC curve analysis plots the proportion of individuals correctly classified among those who died (true-positives/deaths = sensitivity) versus the proportion of individuals incorrectly classified among those living (false-positives/survivors = 1 − specificity) for various cut points of the predictor score, i.e., predicted mortality risk or age difference. A center line is called the "chance line" because, for every point, the probability of a true-positive response equals the probability of a false-positive response, i.e., does not predict any better than chance. The greater the area of the curve above this chance line, the greater the predictive accuracy of the instrument (HRA program). Points below the diagonal reflect values of false-positive greater than those of true-positive, i.e, worse results than would be expected by chance alone. Statistical tests were conducted using the Corroc2 program[17] in order to compare the ROC curves of the CDC and Carter Center HRA programs.

TABLE 1—Number (n) of Subjects, Predicted Risk (HRA)[a], Observed 10-Year Mortality Risk (R), and Age-Adjusted 10-Year Mortality Risk (aR)[b] by HRA Method and Difference between Actual Age and Risk Age for Men

Actual Age Minus Risk Age	CDC HRA				Carter Center HRA			
	n	HRA, %	R, %	aR, %	n	HRA, %	R, %	aR, %
>+1[c]	37	1.40	0.0	0.0	464	6.62	8.8	8.0
−1 to +1	212	2.46	0.5	0.5	597	4.20	4.2	3.7
−5 to −1	512	6.00	2.5	2.4	363	6.75	9.6	9.0
−10 to −5	489	16.10	12.3	11.6	62	12.30	21.0	20.3
<−10	257	29.37	17.1	16.4	21	13.32	19.0	17.9
Overall	1507	12.65	7.8	7.4	1507	6.02	7.8	7.2

[a] Average HRA-predicted 10-year mortality risk.

[b] Age-adjusted risks for each sex and HRA method are predicted from the results of four logistic regression models, where each category of risk is transformed into dummy variables and the average risk (−1 to +1) category is left out of the model; age is set equal to 39.6 years (the total sample mean).

[c] Expected to be lowest risk category.

TABLE 2—Number (N) of Subjects, Predicted Risk (HRA)[a], Observed 10-Year Mortality Risk (R), and Age-Adjusted 10-Year Mortality Risk (aR)[b] by HRA Method and Difference between Actual Age and Risk Age for Women

Actual Age Minus Risk Age	CDC HRA				Carter Center HRA			
	n	HRA, %	R, %	aR, %	n	HRA, %	R, %	aR, %
>+1[c]	33	.64	0.0	0.0	84	4.05	1.2	5.0
−1 to +1	576	1.52	0.9	0.9	672	1.98	1.6	2.4
−5 to −1	673	4.45	3.0	2.9	695	4.75	3.0	3.6
−10 to −5	251	10.60	5.2	5.1	126	10.60	8.7	8.4
<−10	95	22.71	10.5	10.4	51	19.84	7.8	7.2
Overall	1628	5.35	2.9	2.9	1628	4.45	2.9	3.6

[a] Average HRA-predicted 10-year mortality risk.

[b] Age-adjusted risks for each sex and HRA method are predicted from the results of four logistic regression models, where each category of risk is transformed into dummy variables and the average risk (−1 to +1) category is left out of the model; age is set equal to 39.6 years (the total sample mean).

[c] Expected to be lowest risk category.

Results

All 3135 never- or current-smoking, 25 to 60 year olds in the 1959 to 1960 Tecumseh Community Health Survey included in this study were White, and 48% were male. Observed mortality risk was 5.3% after 10 years and 13.9% after 20 years.

Correlation of HRA Predictions and Comparison of Difference between Actual Age and Risk Age

Predicted risks of mortality from the two programs were highly correlated (Pearson's r = .81; 95% confidence interval (CI) = 0.78, 0.84), as were the risk ages (Pearson's r = .92; 95% CI = 0.88, 0.95). These correlations remained high when examined by age group and sex. However, the difference between risk age and actual age (age difference) for the two methods was not correlated (Pearson's r = .18; 95% CI = 0.14, 0.21). The average age differences were −0.8 years for the Carter Center HRA and −4.2 years for the CDC HRA. The average age differences between the two programs were 3.4 years (95% CI = 3.2, 3.6) overall, 5.9 years (95% CI = 5.6, 6.2) for men, and 1.1 years (95% CI = 0.79, 1.4) for women. This difference increased when stratified by actual age.

Comparison of 10-Year Mortality Risks by Difference between Actual Age and Risk Age

A risk age that is the same (−1 to +1 category) as actual age signifies that the individual is at an average risk level for his/her age, race, and sex for the CDC HRA (and for Carter Center HRA presence of nonmodifiable risk factors). Thus, we would expect to see an increase in observed mortality risk as the difference between actual age and risk age increases (from > +1 category to < −10 category). The predicted, observed, and age-adjusted mortality risks by age difference are shown in Tables 1 and 2. Overall, the CDC HRA program classifies more people in the higher risk categories than the Carter Center HRA, and thus the predicted mortality risks are almost double those observed among men (12.65% predicted vs 7.8% observed) and among women (5.35% predicted vs 2.9% observed). In contrast, the Carter Center slightly underestimates these risks among men (6.02% predicted vs 7.8% observed) but overestimates risks among women (4.45% predicted vs 2.9% observed).

With the CDC HRA program, as the difference between actual age and CDC HRA risk age increases for both men and women, so do the predicted risks and the total observed and age-adjusted 10-year mortality risks. Using the Carter Center HRA, the lowest age difference (> +1 category) had a higher predicted and observed mortality risk than the next highest category (−1 to +1), after which risks rise steadily for both men and women (Tables 1 and 2).

Relative Operating Characteristic Curve Analysis

We examined ROC curves for both the CDC and Carter Center HRA-predicted risks by 10-year mortality rates, since HRA predictions are based on the

10-year appraised chance of dying. The ROC curves for both programs were quite similar and were not significantly different from each other for men (P = .925) or women (P = .9775). The HRA predictions by age group for each sex were also examined and showed similar results; predictions were best for those under age 45.

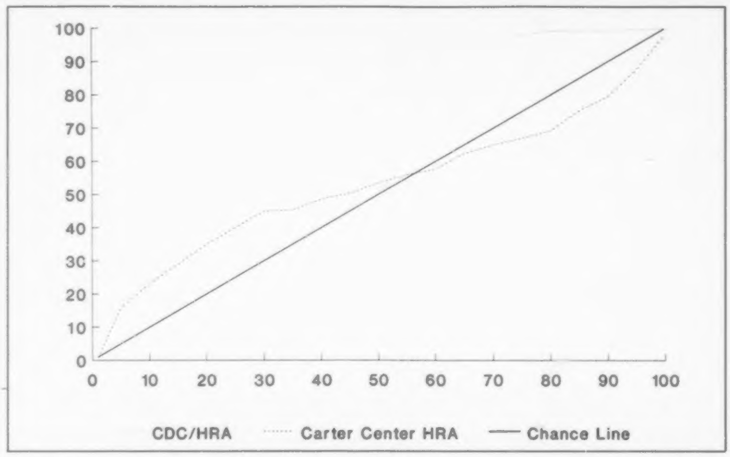

FIGURE 1—ROC Curve Analysis of the Difference between Actual Age and HRA-Predicted Risk Age: Comparison of CDC and Carter Center Methods for *Men*, Using Observed 10-Year Mortality as the Outcome Variable, Tecumseh Community Health Study (n = 1507). Risks were ordered from high to low: CDC difference ranged from −21.5 to 3.7. Carter difference ranged from −16.3 to 7.

The difference between actual age and risk age examined by ROC curves showed that for both men and women (Figures 1 and 2) and within each age group (not shown), actual age minus risk age is a better predictor of mortality with the CDC HRA than the Carter Center HRA. In fact, the ROC curve for the Carter Center HRA was below the chance line for men (Figure 1) but not for women (Figure 2). Thus, for men the Carter Center age difference, on average, predicted 10-year mortality no better than chance alone. As depicted in Figures 1 and 2, these ROC curves were different between the two programs for men (P = .0083) and for women (P = .0257). Further examination of the data in Figure 1 showed that at the point where the Carter Center HRA crosses the chance line (approximately 60%), the difference between actual age and risk age is −3.6 years, whereas for the CDC HRA it is 1.4 years. Moreover, the mean age for those individuals below the chance line (Carter Center) is 43.3 years, and for the same area of the CDC calculations it is 32.7 years. Those individuals below the chance line for the Carter Center HRA were also more likely to be smokers (99.4%) compared with those before the curve crosses the chance line (79.7%).

We also examined this difference by age groups for the two programs. The results were similar for all ages except for those ages 45 to 60, for whom the Carter Center HRA age difference was consistently below the chance line, indicating that it predicted worse than chance alone for this age group (P = .0005).

Discussion

Although mortality risk predictions from the two HRA programs were highly correlated, the calculated risk ages were quite different. It seems unlikely that this result is due to the different reference populations used (the Carter Center program used 1980 to 1982 vital statistics as the reference population and the CDC HRA used 1975 to 1977 vital statistics), because mortality rates have changed little during that time period. More likely, there are differences in how risk age is determined, e.g., mortality risks predicted by the Carter Center more closely approximated those observed than the CDC HRA mortality risks. Further analysis of the specific calculations of risk age is recommended, particularly the implications of using population risk (CDC HRA) versus "adjusted" population risk (Carter Center HRA).

On average, the Carter Center HRA age difference was less than that of the CDC HRA age difference. However, when the age differences for the two programs were compared with the observed 10-year mortality using ROC curve analysis, the CDC HRA age difference was consistently a more accurate (i.e., more sensitive and specific) predictor of individual mortality than was the Carter Center HRA age difference, especially in older men. In addition to performing poorly with the calculation of risk age for older

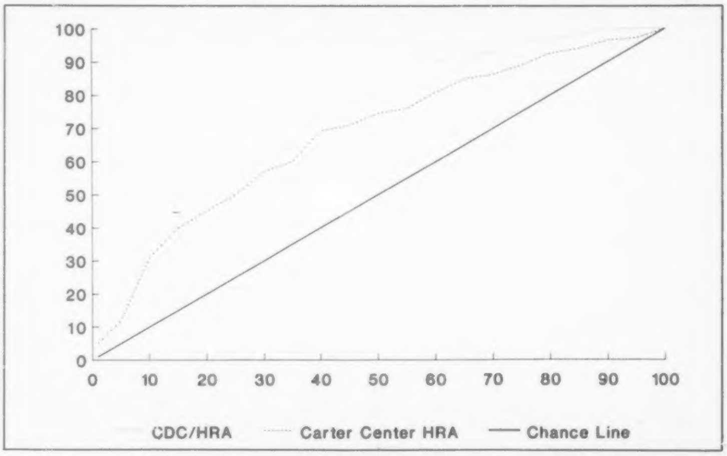

FIGURE 2—ROC Curve Analysis of the Difference between Actual Age and HRA-Predicted Risk Age: Comparison of CDC and Carter Center Methods for *Women*, Using Observed 10-Year Mortality as the Outcome Variable, Tecumseh Community Health Study (n = 1628). Risks were ranked from high to low: CDC difference ranged from −21 to 2.8. Carter difference ranged from −33 to 4.2.

men, the Carter Center HRA may have erred in the weight used for the smoking risk factor.

One of the criticisms of HRA concerns the reliability of self-reported clinical information.[9,10] Reliability was not a problem for blood pressure, weight, and cholesterol in this study, because we used the results of clinical examinations. Although we had data for only approximately 45% of the items used to calculate the predictions (14 of 31 items for CDC/HRA and 13 of 29 items for Carter Center HRA), we did have information on the major risk factors for the common causes of death (e.g., stroke [7.5%], heart disease [39.4%], and lung cancer [5.8%], which accounted for 52.7% of the Tecumseh deaths). Since cardiovascular disease (CVD) is the leading cause of death[18], we also performed ROC curve analysis on the difference between actual CVD risk and predicted CVD risk for each program. The ROC curves for this analysis were virtually identical to those for total mortality; thus the CDC HRA was also a better predictor of CVD mortality.

The variables that we did not have available (e.g., miles traveled per year, social ties) are used as risk factors for less common causes of death, such as motor vehicle accidents and suicide.[8,15] Moreover, we restricted our sample to those individuals older than 25 years of age, and these less common causes of death occur more frequently among younger age groups.[18] Both programs defaulted to population norms for missing values; thus, we would expect the missing items to be extreme to make large changes in the resulting predictions.

Intrinsic errors of the HRA instrument, such as the scientific uncertainty of some risk factors or errors inherent in statistical calculations, will introduce possible problems of accuracy.[15] Our comparison may have been biased toward one program or another because of inherent characteristics of the data used. We attempted to make the comparison as fair as possible by using the same set of variables, calculating ROC curves for 10-year mortality, and using default missing values for each program. Given the constraints of the current software programs and the limitations of the Tecumseh Community Health Study, the CDC HRA appears to be a superior predictor of mortality in the population.

HRA has also been criticized because its effectiveness as a health education/promotion tool has not been demonstrated.[3,8] The revised Carter Center HRA program was primarily developed to update the scientific base of the HRA instrument in order to improve the accuracy of the mortality predictions.[11] If the predictions are improved with this revision, it would seem that this instrument would also be a better health education tool, at least providing more accurate information to the participant. It appears that the revisions did not necessarily improve the mortality predictions (as they were very similar to the CDC HRA mortality predictions) and may have erred on the risk-age calculations for men. If the risk-age calculations are incorrect for a particular program, the effectiveness of that HRA as a health education tool should be questioned. Current HRA users may want to continue using their existing program until the potential problems with the Carter Center Healthier People HRA program are examined further and the effectiveness of the two HRAs in altering behavior has been compared. Those considering the use of the HRA must judge which program better suits the needs of their health education programs. □

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APPENDIX A—CDC vs Carter Center Health Risk Appraisal Program Calculation of an Individual's Overall Predicted Risk of Mortality and Risk Age

CDC HRA	Carter Center HRA
Uses age-sex-race-specific mortality rates for 52 causes of death and "all other" causes	Uses age-sex-specific mortality rates for 41 causes of death and "other causes"
The 10-year expected risk of mortality for each leading cause of death is then "adjusted" by the measured risk factors using the credit-debit method	The individual's current cause-specific mortality risk is estimated using multivariate statistical techniques or population average risk
The cause-specific risks of mortality are summed to determine an overall 10-year expected risk of mortality (predicted risk)	The appraised risks for each cause are summed to get the participant's total appraised risk (predicted risk)
This overall total is compared with age-sex-race-specific risks for the general population to determine "risk age," the age for the given sex-race category with the same overall mortality risk as that predicted from the risk factors	The total appraised risk is compared with either the US population projections for causes of death with modifiable risk factors or an adjusted US population projection for causes of death quantified by nonmodifiable risk factors. This comparison determines the individual's risk age

APPENDIX B—Variables Classified as Precursors and Used Directly by the Carter Center HRA and CDC HRA for Computation of Risk and Those Available from 1959 Tecumseh Community Health Study

Variables Not Available from Tecumseh Study	Variables Available from Tecumseh Study that Required Some Recoding	Variables Available from Tecumseh Study Requiring No Recoding
Smokeless tobacco use[a]	Drinks per week	Age
Miles traveled per year	Hysterectomy	Relative weight
Car size[a]	Diabetes	Blood pressure
Speeding[a]	Cigarettes	Serum cholesterol
Safety belt use	Cigars	Race/origin[b]
Drive after drinking[a]	Pipes	Marital status[b]
Mammogram[a]	Family history of heart disease and diabetes[b]	Sex
HDL cholesterol[a]	Rectal problems[b]	
Pap smear	Lung disease[b]	
Family history of breast cancer		
Menarche[a]		
Age at birth of first child		
Drugs/medication use[b]		
Physical activity level[b]		
Overall health[b]		
Satisfaction with life[b]		
Social ties[b]		
Sleep habits[b]		
Loss/misfortune[b]		
Witness/involved in violent argument[b]		
Risky behaviors[b]		
Breast self-examination[b]		

[a] New variables used by the Carter Center HRA program that were not included in the CDC HRA program calculations.

[b] Variables used by the CDC HRA program that were not included in the Carter Center HRA program calculations.

Impact of a Geriatric Assessment Unit on Subsequent Health Care Charges

William B. Applegate, MD, MPH, Marshall J. Graney, PhD, Stephen T. Miller, MD, and Janet T. Elam, BS

ABSTRACT

Background. In a recent randomized controlled trial assessing the efficacy of an inpatient geriatric assessment unit (GAU), we demonstrated that patients assigned to the unit were less likely than those in the usual care group to be admitted to a nursing home and more likely to remain living in the community. Despite evidence of benefits of GAU care, questions remain about its cost. We examine the impact of GAU treatment on subsequent health care charges.

Methods. Sequential referrals of elderly disabled patients considered at risk for institutionalization or likely to improve with rehabilitation were randomized to the GAU or to usual care. Charge data were collected on utilization of a wide variety of health care services starting at the point of randomization.

Results. The group randomized to the GAU experienced significantly higher rehabilitation charges per participant ($1[illegible] 583 vs $2066, $P = .0001$), but lower mean nursing home charges ($1798 vs $3426, $P = .004$). Total health care charges per participant over the 1 year follow-up were greater for the GAU group ($28 406 vs $16 444, $P = .004$). When charges were adjusted per year of life survived, the GAU group still showed a substantial (but nonsignificant, $P = .10$) increase in total charges. However, when charges were adjusted per day subsequently spent residing in the community, adjusted total charges were similar between the two groups.

Conclusion. Our results indicated that improved outcomes from GAU care require an investment in rehabilitation that is not totally offset by decreased institutional charges in the following year. (*Am J Public Health.* 1991;81:1302–1306)

Introduction

The utility of geriatric assessment has recently been examined by several studies and one National Institute on Aging consensus conference.[1-13] Positive effects of geriatric assessment and subsequent treatment have been demonstrated by Rubenstein and colleagues in a Department of Veterans Affairs Hospital setting and by our group in a community rehabilitation hospital.[14,15] Both of these randomized controlled trials of inpatient geriatric assessment units (GAUs) targeted functionally impaired older patients recovering from acute hospitalization. Both studies demonstrated that GAUs improved patient function, decreased long-term institutionalization, and possibly reduced mortality.[14,15] In our trial, GAU patients were more likely to reside in the community at each follow-up interval.

In our trial, randomization to the GAU or a control group was stratified according to whether the patient was judged prior to randomization to be at higher or lower risk of immediate nursing home placement. When data were analyzed by prerandomization strata, subsequent nursing home days were reduced by GAU treatment for both strata, but physical function was improved and mortality was reduced only in the lower risk stratum.

Despite evidence of benefits of GAU care, questions remain about its cost. In the trial conducted by Rubenstein and colleagues,[14] treatment on the GAU resulted in decreased estimated health care costs per year of life survived. During our study we collected data on both reimbursed and unreimbursed health care utilization charges after randomization and validated the data in an extended data collection project. This article describes the impact of GAU treatment on these charges, both for the entire study cohort and for the two risk strata.

Methods

Patients

Patients who were considered by physicians and social work personnel to be functionally impaired and at risk of nursing home placement were referred to the GAU. Inclusion and exclusion criteria and the screening process have been described previously.[15]

During the 2-year enrollment period 278 referrals were received. Of these, 123 (44%) were determined to be ineligible for the following reasons (percents are of total consultations): medically unstable, 18 (6.5%); terminally ill, 12 (4.3%); too severely cognitively impaired, 34 (12.2%); not sufficiently medically or functionally impaired, 12 (4.3%); refused consent, 21 (7.6%); nursing home placement inevitable, 8 (2.9%); and miscellaneous, 18 (6.5%). The remaining 155 patients were randomized, 77 to the control group and 78 to the GAU group.

Randomization. Randomization was conducted in a stratified block design. Patients were stratified prior to randomization according to an assessment of their risk of immediate nursing home placement at hospital discharge. This assessment

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All authors are with the Department of Preventive Medicine, and Dr. Graney is also with the Department of Biostatistics and Epidemiology, University of Tennessee, Memphis.

Requests for reprints should be sent to William B. Applegate, MD, 66 N Pauline, #232, Memphis TN 38163.

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was based on patient and physician estimate of likelihood of nursing home placement and identified two strata with different levels of risk for subsequent nursing home placement.[15] The methodology for stratification and randomization has been published previously.[15]

The Geriatric Assessment Unit. The GAU studied was a 10-bed geriatric unit based in a rehabilitation hospital that is a separate building in the Baptist Memorial Hospital complex. The organization, specific staff roles, and function of this multidisciplinary unit have been described previously.[2,16] The GAU provided equal emphasis on interdisciplinary assessment of patient problems and on rehabilitation. Particular attention was paid to problems common in frail hospitalized older patients, including incontinence, decubitus ulcers, contractures, confusion, sensory disorders, depression, medication toxicity, inadequate nutrient intake, weakness or lack of motivation, inability to perform activities of daily living (ADL), gait disorders, falls, and degree of family and social support.

After completion of the initial GAU assessments, the first of a series of weekly team meetings was held to determine whether or not the patient was a candidate for a specific treatment and/or rehabilitation. Patients needing medical care were either treated on the GAU or returned to the care of the original referring physician. A rehabilitation plan with specific goals was developed for patients needing rehabilitative care, and patient progress was reevaluated weekly. After discharge from the GAU, patients were returned to their regular physicians and received no further specialized intervention.

Control Group. GAU or research staff made no arrangements for postrandomization care of control group patients. Instead, control group patients received services that they or their physician chose from among options available to any patient, but excluding treatment on the GAU.

Measures

Patient evaluations were made prior to randomization and at 6 weeks, 6 months, and 1 year after randomization. Postdischarge interviews were conducted in the patient's place of residence. All data were collected by trained interviewers who did not provide clinical care but who could not realistically be blinded to group assignment. Data collected included demographic variables, place of residence, self-reported physical activities ranging from basic ADLs to ability to move heavy objects,[17] observed timed physical performance,[18] the CES-D Depression Scale,[19] the Folstein Mini-Mental State Exam,[20] and data on charges.

Charge data were collected on subsequent hospital utilization, physician services, nursing home utilization, inpatient and outpatient rehabilitation, home health care, homemaker/sitter care, medication charges, equipment charges, and other miscellaneous health care-related charges. At the time of randomization, each patient was given a charge data notebook that contained a brief description of the study, a list of possible charges that might be incurred, a chart to record charges, and pockets to store bills, receipts, etc. At the 6-month interview these notebooks were reviewed and, if necessary, patients were reinstructed in charge documentation. At the 1-year interview the notebooks were collected, reviewed with the patient or a family member, and any additional undocumented charges and their sources were noted.

We also constructed a simplified form in layman's language that listed possible categories of health care charges (e.g., hospitalizations, nursing homes, home health care, equipment), asked for the source of these charges (e.g., name of institution, pharmacy, health care agency), and included some open-ended questions such as "Have you been in a nursing home or had someone to care for you in your home?" In cases where a family member or other financial advisor was the responsible party, we mailed the simplified form along with a letter of explanation and a copy of the signed consent form. Follow-up telephone calls were often necessary to verify and explicate charges listed on forms. Telephone calls were also made when forms were not returned, and data were collected from a family member if necessary.

Charge data for all hospital stays, rehabilitation stays, skilled care nursing stays, and home health care were verified with the sources of care. All cooperated, including public and private institutions throughout the United States. Also, length of stay and daily rate were verified with source of care for nursing homes and board and care homes.

For physician, dentist, optometrist, pharmacy, and equipment charges, bills or receipts were used if available. If not, charges were obtained from the source of care. In the few cases where ledger cards of deceased patients had been destroyed, estimates of physician charges were made from the patient's medical record. Medicine charges for which documentation was unavailable were estimated by the patient or family and this estimate was cross-checked with the 1986 *Pharmacy Redbook.*[21] Physician charge data were cross-checked with pharmacy documents to identify otherwise unreported charges.

Because private home care is overlooked in many studies of health care charges, a special attempt was made to collect charge data for private duty nurses, sitters, housekeepers, and paid companions who were hired because of the patient's medical condition. Patients with substantial home care charges generally had a responsible family member who could provide charge data. Also, a large home care registry cooperated fully in providing charge information.

Valid charge data were collected on 152 of 155 randomized patients. Control group data were lacking for one patient who withdrew at 6 weeks and would not allow further data collection and two patients who could not cooperate due to mental status deficits and whose families refused to cooperate.

Statistical Analysis

Equivalence of study groups was tested using Student's t test of differences between means of two independent samples (pooled estimate of population variance) for continuous variables or the chi-square test of statistical independence for categorical variables.[22] All analyses used the intention to treat approach. All statistical tests were nondirectional, with $P < .05$ used as the criterion of statistical significance. Unless specified otherwise, data reported here on utilization and charges represent group means.

Results

The baseline demographic characteristics and functional status of the 152 patients included in this report were similar to those in the study cohort of 155 reported previously.[15] Characteristics of this study cohort included mean age of 78 years, 78% female, 25% married, and 52% living alone. The cohort exhibited substantial functional impairment with 86% rated either moderately disabled (dependent on 3–6 out of 8 ADLs) or severely impaired (dependent on 7 or all ADLs).

There were no significant differences at the time of randomization between the GAU group and the control group in base-

TABLE 1—Absolute Dollar Charges and Charges per Year Survived[a]

	Absolute Dollar Charges			Charges per Year Survived		
Source of Charges	GAU[b]	Control	P[c]	GAU[b]	Control	P[c]
Rehabilitation units[d]	10 583 ± 6302	2066 ± 4654	.0001	18 406 ± 26 256	2215 ± 4795	.0001
Initial hospitalization[d]	260 ± 1025	1402 ± 3333	.004	279 ± 1081	3148 ± 8868	.005
Subsequent hospitalization[d]	10 401 ± 25 062	5833 ± 8967	.14	33 528 ± 148 157	12 145 ± 23 232	.22
Nursing home	1798 ± 4856	3426 ± 5492	.05	2292 ± 6223	4866 ± 7952	.03
Emergency room, outpatient physician, laboratory	348 ± 563	416 ± 770	.53	392 ± 607	806 ± 2702	.19
Prescription medications	515 ± 443	609 ± 627	.28	539 ± 450	731 ± 765	.06
Home health, housekeeper	3941 ± 10 777	2192 ± 5230	.21	4200 ± 11 383	2744 ± 5729	.32
Other[e]	560 ± 712	501 ± 763	.62	657 ± 862	1478 ± 563	.21
Total	28 406 ± 31 674	16 444 ± 14 504	.004	60 292 ± 159 312	28 134 ± 31 969	.09

[a] Excluding 3 patients with missing data, N=152 (78 GAU, 74 control); data are expressed as mean ± standard deviation.
[b] GAU=geriatric assessment unit.
[c] Two-tailed *t* tests for unpaired data.
[d] Postrandomization charges only; includes physician charges.
[e] Equipment, supplies, nonphysician health professionals, and health-related transportation.

line demographic data, scores on the CES-D Depression Inventory, the Mini-Mental State Exam, ADLs, measures of timed physical performance, or types of diagnoses.

Analysis of baseline characteristics of the strata at higher vs lower risk for immediate nursing home placement revealed that the higher risk strata had significantly longer length of initial hospital stay and higher levels of dependence in ADLs at baseline. Age, race, sex, marital status, and prior residence did not differ between the two strata.

Table 1 compares the charges for the GAU vs control groups during the 12 months of follow-up after randomization. The GAU group had significantly higher charges for inpatient rehabilitation (P = .0001) and absolute total charges over the year of follow-up (P = .004). Because admission to the GAU allowed rapid placement, the GAU group had significantly lower hospital charges immediately after randomization (P = .004). The GAU group also incurred significantly lower nursing home charges (P = .05). However, in computing total absolute charges, these lower nursing home and initial hospital charges were not of sufficient magnitude to offset the GAU group's higher rehabilitation charges. As can be seen in Table 1, there were no statistically significant differences between the groups for any other category of absolute charges.

Analysis of the charge data by risk strata (not shown in Table 1) indicated that absolute total charges for the GAU group in the higher risk stratum were significantly higher than those for the corresponding control group for 1 year after randomization ($38 793 vs $16 514 per participant, P = .02). As expected, rehabilitation charges were significantly higher for the high-risk GAU group than for the high-risk control group ($13 118 vs $1888, P = .0001). In the higher risk stratum, mean subsequent nursing home charges tended to be lower for the GAU group ($3078 vs $4826, P = .28) over the year of follow-up, but these differences were not significant and did not offset the overall higher rehabilitation charges. In the lower risk stratum, the GAU group had significantly higher rehabilitation charges compared to the control group ($9164 vs $2168, P = .0001) and the GAU group showed a trend toward higher total charges during the 12-month follow-up period ($22 590 vs $16 404, P = .06). Again, the GAU group in this stratum showed a trend toward lower subsequent nursing home charges ($1081 vs $2621, P = .10), but this trend was not enough to offset overall absolute charges.

Charges were also adjusted per year of life survived to account for persons who died (Table 1). Even when adjusted per year of life survived, the GAU group had significantly higher mean charges than the control group for inpatient rehabilitation ($18 406 vs $2215, respectively; P = .0001). The GAU group had significantly lower adjusted hospital charges immediately after randomization ($279 vs $3148, P = .005) and nursing home charges ($2292 vs $4866, P = .03). The GAU group also demonstrated a trend toward lower medication charges ($539 vs $731, P = .06). However, total adjusted (per year of life survived) charges across all categories were still higher for the GAU group ($60 292 vs $28 134, P = .09), but this difference did not reach statistical significance. However when total charges were adjusted per subsequent day of life spent in the community, the total charges were similar for the two groups ($246 vs $331 per participant).

Discussion

We previously reported that a randomized clinical trial of a geriatric rehabilitation unit in a community rehabilitation hospital had favorable results in allowing elderly persons to continue their residence in a home and demonstrating improved physical function.[15] This article extends those observations by addressing the question: "What are the financial investments required to achieve these outcomes?" Our results indicate that if broadly defined medical care charges are used as the measure of subsequent economic expenditures, improved outcomes require an investment in rehabilitation that is not recovered during the following year. When total unadjusted charges during the 12 months of follow-up were analyzed, a statistically significant increase in absolute total charges for health services was observed for the GAU group. This increase in charges was produced primarily by rehabilitation charges. The mean charge for subsequent nursing home services in the GAU group was less than the comparison group, reflecting the greater likelihood that a patient rehabilitated in the GAU would remain living in the community. However, the difference in mean charges for nursing home services between GAU and comparison groups did not fully offset the greater charges for care in the rehabilitation hospital.

Absolute charges provide one economic perspective, but they can be misleading because patients who die do not generate further health care charges. Also, absolute charges may indicate perspectives of payors but ignore the residential and functional status of the patient. To counteract these problems, findings were also analyzed as charges adjusted per year of life survived and per day of life spent in a community setting subsequent to randomization. Total charges per year of life survived were still substantially higher for the GAU group, but did not reach statistical significance after adjustment. When the total charges for the GAU and control groups were adjusted per day of life spent in a community setting, there was neither a significant difference between the two groups in total charges nor a substantial trend indicating some potential difference. Therefore, although absolute total charges for the GAU group were significantly greater in the year subsequent to randomization, this expense was balanced by the fact that GAU patients had fewer nursing home admissions and were more likely to remain in a community setting.

This study obtained data on actual charges rather than estimating costs. Both approaches have strengths and weaknesses. Health care charges do not necessarily reflect actual costs for delivering a service. On the other hand, estimating costs involves judgments that can have substantial impact on the direction of the ultimate results.[23] Few institutions currently have accurate cost data available. We chose to use actual charges rather than estimated costs because charges more clearly reflect the actual expense of GAU and subsequent care to patients or payors, given current reimbursement methods. A potential limitation of our charge data concerns the reliability of the charge data for noninstitutional services. Items such as housekeeper, home health services, and pharmacy charges may be remembered less reliably. However, we did make every effort to document these categories of charges with source of care and the relatively small magnitude of noninstitutional charges is such that some inaccuracy in these data would not alter the findings of the study.

In the trial conducted in a Department of Veterans Affairs Hospital using estimated costs, GAU patients had similar absolute costs when compared to control patients ($19 406 for the GAU vs $20 329 for the control).[14] When these costs were adjusted per year of life survived, the GAU had lower costs than the control group ($22 597 vs $27 826, respectively). By comparison, in our study, GAU care actually resulted in higher total absolute charges over the 1-year study period. Further, when charges were adjusted per year of life survived, the GAU group still demonstrated a trend toward higher total charges over the year of study, although this difference was not statistically significant. These findings are different from those reported by Rubenstein et al.[14] because our GAU had much higher absolute daily charges. However, higher charges in the community hospital GAU in this study may reflect reimbursement strategies rather than the actual costs essential to mount a successful GAU intervention. Also, our GAU group tended to have higher subsequent hospitalization charges, while the opposite was true in the Rubenstein study. However, all of our study participants were returned to their usual source of care after the GAU while participants in the Rubenstein study continued to receive specialized follow-up care. This difference in design might account for our different findings with regard to subsequent hospital utilization.

The cost-effectiveness data presented here demonstrate an important issue for health policy analysis of services provided to frail older persons. Although absolute charges may be higher for a new form of health care delivery, analysis of charge or cost data should include some form of adjustment per unit of benefit received. Our study found that GAU care led to greater total absolute charges and greater charges per year of life survived. However, when charges were adjusted per day subsequently spent residing in the community, total charges between the two groups were similar. ☐

Acknowledgment

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Important Aspects of Right-to-Know Legislation Not Being Used by Communities, Say Tufts Researchers

Although communities can obtain specific information about the manufacture, use, storage, transportation and emission of toxic chemicals within their borders, local planning committees generally have not been using such information to prevent pollution or reduce risks, according to a study done by researchers at Tufts University's Center for Environmental Management.

"An in-depth study of four communities suggests that the right-to-know aspects of the 1986 federal Emergency Planning and Community Right-to-Know Act may be taking a back seat to emergency planning activities," said Kathleen Rest, former project manager and assistant professor of occupational and environmental health at the University of Massachusetts Medical Center. "Such information could be used to persuade local governments and industry to go beyond emergency planning and get involved in accident prevention, risk reduction, and pollution prevention activities," she said.

Congress passed the act, also known as SARA Title III, in 1986, partly in response to the chemical disaster in Bhopal, India, which killed more than 2000 people and injured another 100 000. The law requires communities to establish local emergency planning committees (LEPCs) to develop emergency response plans for such environmental disasters using information previously unavilable to local officials, emergency responders, and citizens.

"So far, LEPCs may be taking a narrow view of their authority and responsibility. These committees could play an important role in prevention, but they will have to be persuaded, and they will need help," Rest said.

The study focused on four communities with different types and levels of industrial activity: Springfield, Mass.; Baytown, Texas; Texas City, Texas; and Newark, New Jersey. The researchers looked at how all local communities involved in the Title III activities obtained, used, interpreted and communicated information about hazardous substances in their communities.

"It's a lost opportunity," said Sheldon Krimsky, principal investigator and professor of urban and environmental policy at Tufts. "The LEPCs could play a more proactive role, but they are going to have to help local citizens understand what the new information means," he said.

The researchers identified specific problems and made recommendations on how to improve the implementation of the law's community right-to-know provisions:

- The US Environmental Protection Agency could organize a task force to discuss realistic roles for the LEPCs and provide guidelines for their involvement in prevention activities.
- LEPCs could be encouraged to be inclusive of all community interests when it appoints community members.
- The LEPCs could share data on accidental releases, chemical inventories, and emissions with public health officials.
- The media could play a significant role in making risk information accessible to the public.
- Grassroots citizen and environmental groups have been successful in motivating citizens to take action about toxic hazards issues. LEPCs should work with these groups to enhance public involvement.
- LEPCs should become forums for constructive interaction between citizens, local government, industry, and the public health community.

"If Congress would provide funding to support state and local community activity, and if states would enact statutes to fund Title III activities through fees or appropriations, then the committees would have some incentive to extend their purview and become more effective advocates for prevention activities," said Krimsky.

Public Health Briefs

Mortality of Intravenous Drug Users in Rome: A Cohort Study

Carlo A. Perucci, MD, Marina Davoli, MD, Elisabetta Rapiti, MD, Damiano D. Abeni, MD, and Francesco Forastiere, MD

ABSTRACT

A historical cohort study was carried out in Rome to examine overall and cause-specific mortality among intravenous drug users (IVDUs). A total of 4200 IVDUs (3411 men and 789 women) enrolled in methadone treatment centers between 1980 and 1988 were studied. There were 239 deaths during the follow-up period. The overall SMR was 10.10 in the entire cohort (95% confidence interval, 8.86–11.47), 9.30 in males and 18.07 in females. A large excess of mortality in both sexes was found for infectious, circulatory, respiratory, and digestive diseases as well as for violence, overdose, AIDS, and unknown or ill-defined causes. Tumors and suicide were excessive only in males. Deaths due to drug overdose, violence or trauma, and cirrhosis accounted for 63.6%, AIDS for 7.1%, endocarditis and other bacterial infections for 7.1%, and neoplasms for 3.8% of total mortality. These findings document serious health consequences of drug abuse in Italy. (*Am J Public Health.* 1991;81: 1307–1310)

Introduction

Since the emergence of the acquired immunodeficiency syndrome (AIDS) epidemic most studies have focused on HIV incidence, prevalence, and risk behaviors among intravenous drug users (IVDUs)[1–5] with less attention devoted to overall morbidity and mortality data in this population. There are some indications, however, that IVDUs are at increased risk of overall mortality, deaths from overdose, trauma, infectious diseases.[6–12] This study describes the pattern of mortality in a large cohort of IVDUs in Rome during the period of transition into the "AIDS era."

Methods

Twenty Public Treatment Centers (PTCs) for drug abuse are operating in Rome (population=3 million) since the 1980s. PTCs are open to drug users who reside in each catchment area and agree to a specific outpatient treatment plan. Short-term detoxification with low-dose methadone is currently offered, but subjects usually undertake several methadone cycles.[13]

The three largest PTCs were selected for the study. Information from their clinical records allowed us to enroll a cohort of all the drug users who presented at least once to one of these three centers during the period 1980–1988.

Vital status as of December 31, 1988 was ascertained through the registry office of the last municipality of residence. Cause of death was ascertained through record linkage with the national mortality file for deaths occurring between 1980 and 1987.[14] For deaths occurring in 1988 or not found at the linkage procedure, death certificates were retrieved from the municipality of death and coded by an experienced nosologist according to the ninth revision of the International Classification of Diseases (ICD-9).

For the calculation of person-years at risk of dying,[15] each subject was considered from the date of first enrollment through the end of 1988 or to the date of death. Standardized mortality ratios (SMRs) and their 95% confidence intervals (CI)[16] were used to compare the mortality experience of drug users with that of the total national population. The expected number of deaths was calculated using sex- and age-specific national mortality rates.

Results

A total of 4473 subjects were initially enrolled; 273 were excluded because the identifying data on their records were insufficient to verify vital status, yielding a final sample of 4200 IVDUs (Table 1).

There were only 23.7 expected deaths compared with the 239 observed (SMR-10.10, 95% CI=8.86–11.47) Tables 2 and 3 show sex- and cause-specific mortality. Excess mortality was found for both sexes in all categories. The excess mortality is mainly due to overdose at 3.38 deaths/1000 person-years for both sexes, but violence (1.62/1000 person years) and

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The authors are with the Epidemiology Unit of the Regional Health Authority, Rome, Italy.

Requests for reprints should be sent to Carlo A. Perucci, Epidemiology Unit, Regional Health Authority, via di Santa Costanza, 53, I-00198 Rome, Italy.

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Editor's Note. See related editorial by Peter A. Selwyn on page 1247 of this issue.

TABLE 1—Selected Characteristics of Intravenous Drug User Study Population, Rome, Italy, 1980–1988

Characteristics	Males (n = 3411) No.	Males (n = 3411) %	Females (n = 789) No.	Females (n = 789) %
Age of entry into PTC[a]				
<20	428	12.5	150	19.1
20–24	1644	48.2	355	45.2
25–29	981	28.7	191	24.3
>29	361	10.6	90	11.4
Marital Status				
Unmarried	1783	64.1	329	52.7
Married	999	35.9	295	47.3
Education level				
Primary school	821	28.7	79	12.4
Junior high school	1404	49.2	283	44.6
High school	600	21.0	261	41.1
University	31	1.1	12	1.9
Employment status				
Employed	1392	49.4	195	29.2
Unemployed	1428	50.6	474	70.8
Year of entry into PTC				
1980–82	2152	63.0	452	57.5
1983–85	668	19.6	174	22.1
1986–88	594	17.4	160	20.4
Age at first drug use				
<15	85	2.8	35	5.1
15–19	1448	47.7	329	47.6
20–24	1130	37.2	237	34.3
>24	371	12.3	90	13.0
Primary IV drug				
Heroin	3264	96.8	743	95.7
Other	107	3.2	33	4.3
Frequency of use				
Daily	2492	73.1	590	74.8
Less than daily	335	9.8	57	7.2
Unknown	584	17.1	142	18.0
Main treatment				
Methadone	2675	78.4	599	75.9
Psychotropic drugs	398	11.7	86	10.9
Morphine	67	2.0	25	3.2
Others and unknown	116	3.4	28	3.5
None	155	4.5	51	6.5
Vital Status				
Alive	3074	90.1	709	89.9
Dead	200	5.9	39	4.9
Unknown[b]	137	4.0	41	5.2

[a] Public Treatment Center.

[b] Subjects lost to follow-up were considered alive as of December 31, 1988.

cirrhosis (0.88/1000 person years) also contribute. One AIDS death was recorded for the first time in 1986 (0.4/1000 person years); the AIDS death rate increased rapidly in 1987 (7 deaths, 1.6/1000 person years) and 1988 (9 deaths, 2.7/1000 person years).

The SMR for all causes is higher among females than among males even though the excess rates are similar (9.1 and 8.5/1000 person years in males and females, respectively). An increased mortality from suicide is seen only among males; on the other hand SMRs for pneumonia and septicemia are higher in females. Finally, the SMR for malignancies is higher only in males.

Discussion

The study shows that risk of death increased tenfold in male and twentyfold in female IVDUs. The excess risk of death is observed for most causes of death, including neoplasms in males.

Our study population may be different both in character and in exposure to toxic substances from street addicts and occasional users, thus limiting the generalizability of the results. Moreover, IVDUs are a dynamic population, with some current users permanently stopping drug use each year, but we could not take these changes into account because no information was available.

National standards rates were employed to increase the stability of reference numbers, especially in the young population. However, a more restricted reference population, i.e., Rome, would have assured comparability. Therefore, we examined the cause-specific SMRs for 15- to 34-year-olds in the city of Rome compared to Italy during the period 1980 to 1984; SMR values approximate unity for all categories except overdose, endocarditis, and violence with SMRs of 3.7, 1.67, and 0.73, respectively. On that basis, only slightly different results could have occurred using local standards.

We could not control for other potential confounders (i.e., social class, smoking, and alcohol consumption) besides age and sex. Part of the increased risk of death could be explained by the fact that the educational level and the occupational status of our cohort members indicate a lower social class level. However, a study performed in Northern Italy showed that the differences in the overall mortality for social class are not greater than twofold.[17]

Our data agree with published findings about an excess mortality for cirrhosis, endocarditis, septicemia, trauma, suicide, noninfectious heart disease, and AIDS.[6–12, 18,19] Unlike a study performed in New York City, we did not find an excess mortality from tuberculosis[19]; on the other hand, an Italian report showed only a marginally increased risk for tuberculosis among seropositive drug injectors.[20] Females seem more prone to death due to infectious agents (pneumonia and sepsis). We have no explanation for such differences or data to support Stoneburner and colleagues' hypothesis that some of these deaths could be due to undiagnosed HIV-related diseases.[12,19]

Deaths from overdose due to opiates are high. Most of them (72 of 81) occurred when subjects had been out of treatment for more than 2 weeks. It should be noted also that overdose deaths are probably underestimated because some of them are incorrectly ascribed to "ill-defined causes of death" or "other heart diseases."

The excess mortality from neoplasms in males is, to our knowledge, unprecedented in the literature. Even though chance alone could be responsible for this result, smoking and alcohol consumption could be confounding factors, given their strong association with illicit drug use.[21,22] Immunosuppression or even HIV infection might also play a role in cancer development, but this hypothesis requires further investigation.

TABLE 2—Cause-specific Mortality among Male Intravenous Drug Users (19 662 person-years)

Cause of Death[a]	OBS	EXP	SMR[b]	95% CI[c]
All causes (000–999)	200	21.50	9.3	8.1–10.7
Septicemia (038)	2	0.30	67.7	8.2–244.8
Tuberculosis (010–018)	1	0.07	—	—
All malignant neoplasms (140–239)	8	2.97	2.7	1.2–5.3
Nasopharynx (147)	1	0.07	—	—
Digestive organs (150–159)	4	0.58	6.9	1.9–17.6
Stomach (151)	1	0.18	—	—
Colon (153)	2	0.16	12.3	1.5–44.4
Pancreas (157)	1	0.07	—	—
Respiratory organs (160–165)	1	0.42	—	—
Brain (191)	1	0.25	—	—
Lymphatic and hematopoietic tissue (200–207)	1	0.19	—	—
AIDS (279)	16	—	—	—
Overdose (304)	65	—	—	—
Diseases of nervous system (320–389)	2	0.60	3.3	0.4–12.0
Diseases of circulatory system (390–459)	22	2.60	8.5	5.3–12.8
Hypertensive disease (401–405)	2	0.06	31.8	3.9–115.0
Ischemic heart disease (410–414)	1	0.95	—	—
Endocarditis (421–424)	6	0.08	77.1	28.3–165.8
Other forms of heart disease (420;422–423;425–429)	10	0.60	16.5	7.9–30.4
Cerebrovascular disease (430–438)	1	0.60	—	—
Diseases of arteries or veins (460–519)	2	0.30	6.7	0.8–24.1
Diseases of respiratory system (460–519)	3	0.48	6.2	1.3–18.1
Pneumonia (480–486)	2	0.22	9.2	1.1–33.2
Diseases of digestive system (520–579)	20	1.21	16.5	10.1–25.6
Cirrhosis (571)	18	0.81	22.2	13.2–35.1
Injuries and poisoning (800–999)	46	11.58	4.0	2.9–5.3
Suicides (E950–959)	11	1.61	6.8	3.4–12.2
Ill-defined conditions (780–799)	6	0.50	12.1	4.4–26.2
Unknown causes	9	—	—	—

Note. OBS = observed deaths; EXP = expected deaths; SMR = standardized mortality ratio; CI = confidence interval.

[a] Code numbers from the International Classification of Diseases, Ninth Revision, 1975.

[b] SMR were not computed for categories with fewer than two deaths.

[c] Two-tailed test.

In conclusion, our findings indicate that drug abuse is an important cause of mortality in the young in Rome as well as in New York City. Preventive and treatment strategies are urgently needed. □

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TABLE 3—Cause-specific Mortality among Female Intravenous Drug Users (4334 person-years)

Cause of Death[a]	OBS	EXP	SMR[b]	95% CI[c]
All causes (000–999)	39	2.16	18.1	12.9–24.7
Septicemia (038)	2	0.00	509.9	62.0–1842.2
Tuberculosis (010–018)	—	0.01	—	—
All malignant neoplasms (140–239)	1	0.61	—	—
Nasopharynx (147)	—	0.01	—	—
Digestive organs (150–159)	—	0.11	—	—
Stomach (151)	—	0.04	—	—
Colon (153)	—	0.03	—	—
Pancreas (157)	—	0.01	—	—
Cervix uteri (180)	1	0.00	—	—
Respiratory organs (160–165)	—	0.03	—	—
Brain (191)	—	0.03	—	—
Lymphatic and hematopoietic tissue (200–207)	—	0.14	—	—
AIDS (279)	1	—	—	—
Overdose (304)	16	—	—	—
Diseases of circulatory system (390–459)	5	0.37	13.5	4.4–31.6
Hypertensive disease (401–405)	—	0.02	—	—
Ischemic heart disease (410–414)	—	0.06	—	—
Endocarditis (421–424)	1	0.01	—	—
Other forms of heart disease (420;422–423;425–429)	2	0.10	20.8	22.1–655.7
Cerebrovascular disease (430–438)	2	0.12	16.1	1.9–58.2
Diseases of arteries or veins (440–459)	—	0.06	—	—
Diseases of respiratory system (460–519)	3	0.08	37.1	7.6–108.5
Pneumonia (480–486)	3	0.03	88.0	18.1–256.2
Diseases of digestive system (520–579)	3	0.12	24.4	5.0–71.3
Cirrhosis (571)	3	0.07	39.7	8.2–116.0
Injuries and poisoning (800–999)	4	0.60	6.7	1.8–17.1
Suicides (E950–959)	—	0.13	—	—
Ill-defined conditions (780–799)	2	0.04	45.5	5.5–164.5
Unknown causes	2	—	—	—

Note. OBS = observed deaths; EXP = expected deaths; SMR = standardized mortality ratio; CI = confidence interval.

[a] Code numbers from the International Classification of Diseases, Ninth Revision, 1975.

[b] SMR were not computed for categories with fewer than two deaths.

[c] Two-tailed test.

Leptospirosis on Kauai: Investigation of a Common Source Waterborne Outbreak

ABSTRACT

During the summer of 1987, a common source waterborne outbreak of leptospirosis occurred on the island of Kauai in the state of Hawaii. Eight leptospirosis cases were identified using the microscopic agglutination test. Methodologic weaknesses in this technique were responsible for the initial missed diagnosis on six of the eight cases. All cases had repeated exposure to the Waimea River. A wide array of animal reservoirs and optimal water conditions may have been factors in this outbreak. *(Am J Public Health.* 1991;81:1310–1312)

Alan R. Katz, MD, MPH, Sally Jo Manea, BSN, and David M. Sasaki, DVM, MPH

Introduction

Since the Hawaii State Department of Health (DOH) began collecting detailed statistics on leptospirosis in the early 1970s, Hawaii has consistently had the highest national annual incidence rate.[1] In 1987, the DOH reported a rate of 2.5 per 100 000 population[2] compared with the national rate of 0.02 per 100 000 population.[3] The following report describes a common source outbreak of leptospirosis from recreational exposure to a fresh water source.

Alan R. Katz is with the University of Hawaii School of Public Health. Sally Jo Manea and David M. Sasaki are with the Hawaii State Department of Health.

Request for reprints should be sent to Alan R. Katz, MD, MPH, Department of Public Health Sciences, University of Hawaii School of Public Health, 1960 East-West Road, Honolulu, HI 96822.

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TABLE 1—Final Microscopic Agglutination Test (MAT) Results

Case No.	Acute Titer	Days after Symptom Onset[a]	Convalescent Titer	Days after Symptom Onset[a]	Presumptive *Leptospira interrogans* Serovar
1	0	21	1:200	62	*australis*
2	1:50	7	1:1600	74	*australis*
3	1:50	4	1:400	61	*australis*[b]
4	1:50	4	1:800	74	*australis*
5	0	3	1:200	25	*bataviae*[c]
6	1:100	3	1:3200	71	*australis*
7	0	0	1:3200	23	*australis*
8	0	1	1:1600	19	*fort bragg*

[a] The number of days after the onset of symptoms each sample was obtained.
[b] Blood culture positive for serovar *bangkok* (Australis serogroup).
[c] Blood culture positive for serovar *bataviae* (Bataviae serogroup).

The Outbreak

On July 14, 1987, the DOH was notified of three youths hospitalized with suspected leptospirosis. All had repeated exposures to the Waimea River via swimming. The river runs within the lush Waimea Valley on southwestern Kauai. The valley and hills above the river contain cattle, feral goats, pigs, mongooses, and other rodents.

In April 1987, several spontaneous abortions occurred among a herd of 50 cattle pastured adjacent to the swimming area. Sera from the affected animals (9/9) tested positive for *Leptospira interrogans* serovar hardjo (Sejroe serogroup), with microscopic agglutination test (MAT) titers of 1:800 or greater (personal communication Dr. Craig Nishimoto, 1988, oral).

Methods

Case Finding

The three case-patients and their parents were interviewed in an attempt to identify others with recent river exposure. Sixteen males were identified and sampled, including the three hospitalized youths, who swam at this site almost daily between June 14 and July 4, 1987. Sera were sent to the Leptospirosis Reference Laboratory at the US Centers for Disease Control for MAT evaluation.

A "confirmed" case was defined as an individual with recreational exposure to the Waimea River during the abovementioned dates, and positive leptospiral culture or serologic evidence of recent leptospiral infection manifested by a minimal fourfold increase in MAT titers in paired acute and convalescent sera.

Environmental Investigation

Rodents were trapped at and above the exposure site. Water and mud samples downstream from the swimming spot were cultured for pathogenic leptospires. Seroprevalence data from an island-wide leptospirosis cattle survey conducted in late 1987 by the Hawaii State Department of Agriculture (DOA) were evaluated. Ongoing US Geological Survey measurements on the Waimea River near the exposure site were also evaluated.

Results

Eight individuals reported a flu-like illness with onset between July 8 and July 21, 1987. Symptoms included headache 100% (8/8), fever 100% (8/8), chills 87.5% (7/8), nausea 87.5% (7/8), vomiting 75% (6/8), myalgia 62.5% (5/8), diarrhea 50% (4/8), and jaundice 12.5% (1/8). Five of the eight were hospitalized (62.5%), two were treated as outpatients, and one did not seek medical consultation.

Initial MAT paired sera were positive for only two of the eight symptomatic individuals: one demonstrated *L. interrogans* serovar *bataviae* (Bataviae serogroup), and one demonstrated *L. interrogans* serovar *fort bragg* (Autumnalis serogroup). After blood cultures from an individual with a negative MAT grew *L. interrogans* serovar *bangkok* (Australis serogroup), all sera were retested with a new battery of antigens including serovar *australis* (Australis serogroup), which had been missing from the initial antigen battery. The six with initial negative MAT results demonstrated *L. interrogans* serovar *australis* in the retest (Table 1).

Two cases demonstrated positive blood cultures: one with serovar *bangkok* (as noted above) and one with serovar *bataviae*. The eight asymptomatic males were negative serologically for leptospirosis.

Cases ranged in age from 10 to 24 years (median age=14). The duration of symptoms ranged from 2 to 7 days, with a median of 6 days.

Kidney cultures from 14 rodents trapped near the exposure site as well as water and mud samples from the Waimea River were all culture negative for pathogenic leptospires.

The median river flow rate near the exposure site during the outbreak exposure period was 0.31 m^3/second. The 1987 annual mean flow rate at this site was 1.73 m^3/second. On the two dates most closely bracketing the exposure period, the water temperatures were 22°C and 24°C and the pH measurements were 7.10 and 8.20.[4]

The island-wide DOA cattle seroprevalence survey revealed 89 (64%) of 139 cattle positive for *Leptospira* as manifested by MAT titers of 1:100 or greater against one or more serovars. Of 19 cattle bled on the ranch adjacent to the swimming site, 9 were positive: 8 demonstrated serovar *hardjo*, and 1 demonstrated serovar *bataviae*. Four animals geographically removed from the outbreak demonstrated serovar *australis* (personal communication DOA laboratory report, 1988, written).

Intervention

The local radio station, news media, and press reported the outbreak and warned of the potential danger from swimming in the Waimea River. DOH personnel were on site July 15 and 16 to post warning signs and inform swimmers of the potential danger. In addition individuals obtaining camping or conservation permits were given information on leptospirosis. No subsequent leptospirosis cases were reported from the Waimea River, and only one additional case was reported from the entire island of Kauai during the remainder of 1987.

Discussion

The majority of reported leptospirosis cases both nationally and in Hawaii have been sporadic and unrelated.[5,6] However, cluster outbreaks are reported infrequently.[7-10]

The antecedent epizootic of leptospirosis and the high cattle seroprevalence rate would appear to incriminate cattle as the source of this outbreak. However, because serovar *australis* was not identified in cattle pastured adjacent to the exposure site and because the area contains many potential animal reservoirs, this outbreak may be unrelated to cattle. The small number of feral rodents sampled does not justify exclusion from the analysis.

Optimal water conditions for the survivability of waterborne leptospires—a neutral or slightly alkaline pH[11,12] and temperature of 22°C or higher[13]—were documented in the Waimea River around the outbreak exposure period. Decreased flow rates may have allowed concentration of waterborne *Leptospira*.[10,14] Although river water and mud samples were culture negative, the isolation of pathogenic *Leptospira* by direct culture from soil and natural waters is not routinely done and is generally unsuccessful.[15,16]

Although the MAT is considered a reference test for serological diagnosis of leptospirosis,[17] a negative result does not rule out infection because the patient may be infected with a serotype absent from the battery of testing antigens.[17] This phenomenon was responsible for the initial false negative results on six of eight cases in the present outbreak because serovar *australis* was absent from the testing battery. In addition, the MAT is only partially serovar specific. Cross-reactivity is well documented between serovars in the same serogroup as well as related serovars from different serogroups.[17] This explains the confirmed case with positive blood cultures for serovar *bangkok* but MAT results showing the presumptive presence of serovar *australis*.

Through this experience, we learned that sera from individuals infected with serovar *australis* may not cross-react with antigens from the other serogroups and serovars routinely used in the MAT. To optimize the MAT as a diagnostic tool, the testing battery should include antigens from all leptospiral serovars known to be geographically enzootic. ☐

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Timeliness of National Reporting of Communicable Diseases: The Experience of the National Electronic Telecommunications System for Surveillance

Guthrie Birkhead, MD, MPH, Terence L. Chorba, MD, MPH, Susan Root, MD, Douglas N. Klaucke, MD, MPH, and Norma J. Gibbs

ABSTRACT

The timeliness of reporting four nationally notifiable diseases was examined using data reported via the National Electronic Telecommunications System for Surveillance. Timeliness of reporting varied by disease (bacterial meningitis: median 20 days; salmonellosis: median 22 days; shigellosis: median 23 days; and hepatitis A: median 33 days) and by state. These findings indicate a need to standardize surveillance definitions and to account for reporting differences between states in interpreting regional disease trends or detecting multistate disease outbreaks. (*Am J Public Health.* 1991;81:1313–1315)

Introduction

The timely reporting of communicable diseases to public health authorities is an important part of public health surveillance.[1,2] Prompt reporting permits rapid case investigation and early detection of epidemics. Timeliness is one of the key attributes by which public health surveillance systems should be judged.[3] However, few studies have examined this factor and its potential impact on disease prevention.[4-7]

We determined the delay in reporting of selected communicable diseases using data transmitted from state health departments (SHDs) via the National Electronic Telecommunications System for Surveillance for publication in the *Morbidity and Mortality Weekly Report.*[8] Our findings have implications both for the interpretation of surveillance data and for the design of surveillance systems.

Methods

The Epidemiologic Surveillance Project (ESP) was initiated by the Centers for Disease Control (CDC) in six states in 1984.[8] The system is now referred to as the National Electronic Telecommunications System for Surveillance (NETSS)[1] and receives weekly reports of 49 nationally notifiable diseases from all state and territorial health departments. For this study, NETSS reports were examined for 1987 when 35 SHDs and the District of Columbia participated in the system. The information available for each report included disease, age, race and ethnicity, sex, and the dates of disease onset and report to CDC.

The reporting delay was calculated as the difference in days between date of disease onset and data of report to CDC. Reporting delay was examined for salmonellosis, shigellosis, hepatitis A, and bacterial meningitis (predominantly due to *Neisseria meningitidis* and *Haemophilus influenzae* type B). These diseases were chosen because each may occur in epidemic form so each case report warrants a prompt investigation, and because the diagnosis of each is confirmed by laboratory findings. Reports were analyzed only for states providing the date of disease onset for at least 70% of NETSS cases (Table 1).

Results

The delay in reporting varied markedly by type of disease (Table 1 and Figure 1). The median reporting time ranged from 20 days for bacterial meningitis to 33 days for hepatitis A. More than 75% of reports were made within 8 weeks of disease onset for all diseases. Timeliness of reporting did not differ by the age and race/ethnicity of the patient or by month of disease onset (data not shown). Median reporting delays for salmonellosis and shigellosis varied widely by state, from as little as 4 days for one state to over 8 weeks for another (Figure 2a). The distribution of reporting delays in two states with different patterns in reporting is illustrated in Figure 2b.

Discussion

A principal purpose of public health surveillance of infectious diseases is to prompt public health action in response to changes in disease occurrence.[2] The time-

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At the time of this study, all authors were with the Division of Surveillance and Epidemiologic Studies, Epidemiology Program Office, Centers for Disease Control, Atlanta, Ga. Dr. Birkhead is now with the New York State Department of Health, Albany, NY.

Requests for reprints should be sent to Dr. Guthrie Birkhead, Bureau of Communicable Disease Control, New York State Department of Health, Room 651 ESP, Corning Tower, Albany, NY 12237.

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TABLE 1—Completeness of Coding of Date of Disease Onset for Selected Diseases, National Electronic Telecommunications System for Surveillance (NETSS) 1987 Disease Reports

	Total NETSS Reports			Selected States with ≥70% Onset Date Coding			
Disease	Number of States Reporting	Total Cases	Percent with Date of Onset Coded	Number of States	Total Cases	Percent with Date of Onset Coded	Median Reporting Delay (days)[a]
Shigellosis	35	9771	78	23	7167	92	23
Salmonellosis	35	25 880	72	22	17 624	92	22
Hepatitis A	35	13 758	74	25	9592	90	33
Bacterial meningitis	33	1497	84	28	1241	94	20

[a] Median time in days from date of disease onset to date of disease report to CDC, for cases with date of onset coded in 1987.

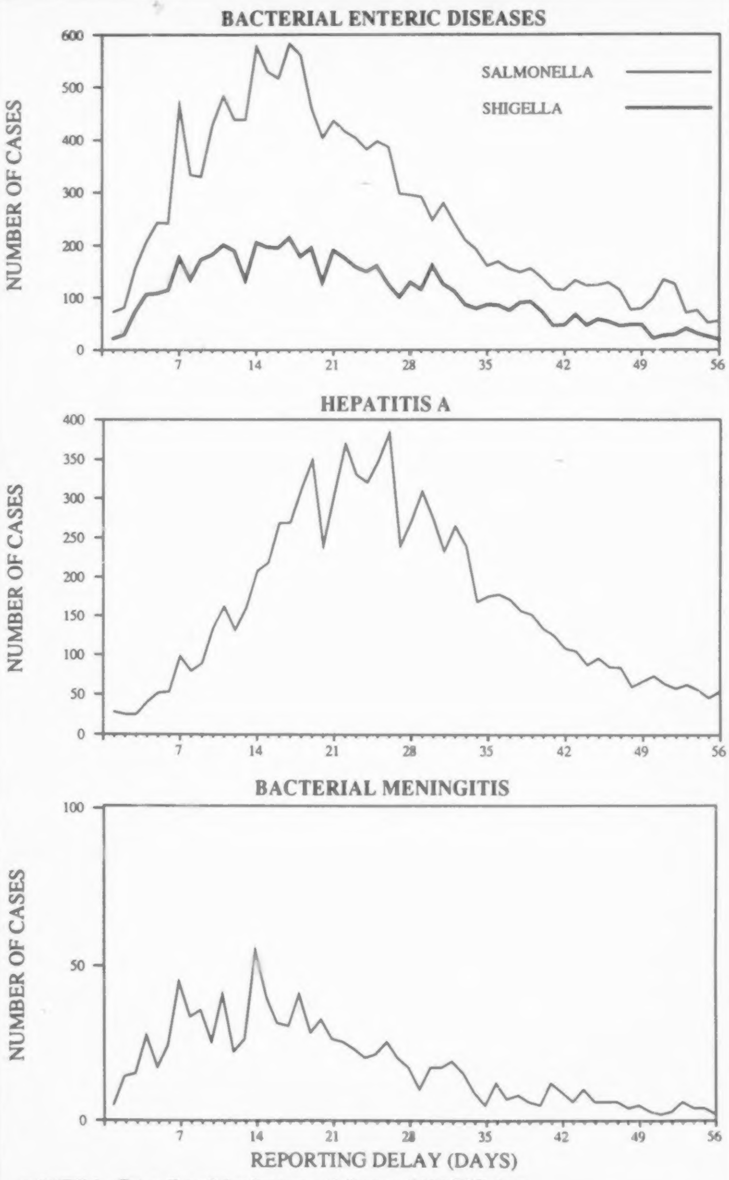

FIGURE 1—Reporting delay by type of disease 1987 ESP data.

liness of reporting is a key element in ensuring that this response is rapid enough to prevent further disease transmission.[1] However, timeliness of reporting and its impact on disease prevention have received little attention in the medical literature.[4–7] We found that the timeliness of reporting to the NETSS system varies widely by type of disease and by state. However, the data available do not allow an evaluation of whether reporting is rapid enough to permit effective disease prevention because timeliness was measured by the report of disease to CDC and immediate preventive action usually takes place at a more local level.

The variability in reporting delay among states may be due to differences among state surveillance systems, including local vs centralized disease reporting, differing statutory requirements for laboratory-based reporting or penalties for failure to report, and differing definitions of disease and of dates of disease onset and disease report. For example, delays in reporting to CDC may be inherent in states in which disease reports go first to county-level health departments before reaching the SHD.

Review of national data does permit detection of regional or multistate disease outbreaks that may be overlooked at the local level. An example of this is the recent recognition of a regional epidemic of *Salmonella* serotype *enteritidis* in the northeastern United States.[9] However, differences in reporting delays between states need to be considered in such an analysis.

Acknowledging the differences in reporting delays by state, the pattern of reporting delays among diseases suggests some factors that may be important in determining reporting time. Reporting of bacterial meningitis was more rapid than for the other diseases studied. This may be due in part to the recognition by physicians and others of the necessity to offer

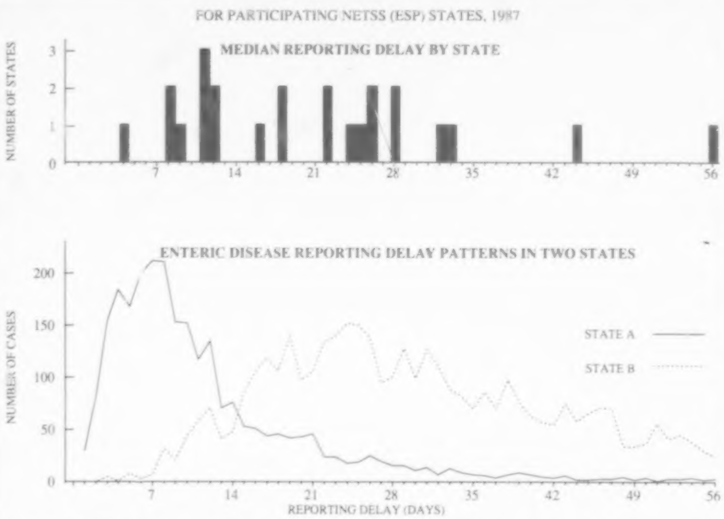

FIGURE 2—Reporting delays from disease onset to report to CDC for enteric disease, by state, for participating National Electronic Telecommunications System for Surveillance States in 1987;
a) = median reporting delay by state;
b) = distribution of reporting delays in two states.

prompt prophylactic treatment with rifampin to close contacts of patients with bacterial meningitis.[10,11] In addition, bacterial meningitis and enteric diseases usually have a more sudden onset and medical care may be sought sooner in the course of disease, allowing more rapid reporting than hepatitis A, which may have a more insidious onset. Finally, bacterial meningitis may be diagnosed from cerebrospinal fluid examination when the patient is first seen by a physician, whereas fecal culture for enteric pathogens and serologic tests for hepatitis A may take several days to complete. This study cannot determine which of these or other factors are most important in determining reporting delay, in part because of the long reporting trail from the physician's office to CDC.

Several areas for improvement in national reporting of surveillance data are highlighted by this analysis. Definitions of date of disease onset and disease report should be standardized. This would parallel recent efforts to standardize disease case definitions and other surveillance system attributes.[3,12,13] For notifiable disease reporting, the date of disease onset is the date of first symptoms attributed to the disease. Date of disease report should be the date of notification of the first-line public health authority responsible for taking immediate preventive action be it town, city, county, or state health department. Health departments should collect the date of disease onset for each reported case. Further studies of the determinants of timely reporting for specific diseases are needed at the local level.

Studies in jurisdictions with unusually short or long reporting times may identify the determinants of the reporting delay. Accounting for inevitable differences in reporting among states in the regional analysis of surveillance data is necessary to take full advantage of national surveillance data for detection of regional or multistate outbreaks. □

Acknowledgments

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The Association between Congenital Syphilis and Cocaine/Crack Use in New York City: A Case-Control Study

Michele S. Zweig Greenberg, MPH, Tejinder Singh, PhD, Mark Htoo, MD, and Stephen Schultz, MD

ABSTRACT

From 1987 to 1989, an epidemic of congenital syphilis was observed in New York City. A case-control study was done to assess the association between various maternal risk factors and congenital syphilis. Independent of the effect of other factors, the odds of being exposed to cocaine were 3.9 times greater among cases than controls (95% confidence interval, 2.8–5.3). This study suggests that the epidemic of congenital syphilis may be related to the increase in cocaine/crack use among delivering mothers. (*Am J Public Health.* 1991;81:1316–1318)

Introduction

The current epidemic of primary and secondary syphilis among women in New York City (NYC) has been characterized by a doubling in the number of cases between 1985 and 1989[1,2] and has contributed to a secondary epidemic of congenital syphilis. Although fewer than 50 congenital syphilis cases per year were reported to the NYC Department of Health (DOH) between 1980 and 1986, 1017 cases were reported in 1989.[2] During this same period, the proportion of delivering mothers in NYC who reported use of cocaine/crack during pregnancy increased dramatically from 1.0 per 1000 live births to 23 per 1000 live births[3] (Figure 1). Because an association between sexually transmitted diseases and cocaine/crack use had been suggested previously,[4–10] we conducted a case-control study to assess the relationship between various maternal risk factors and congenital syphilis in NYC.

Methods

Congenital syphilis is a reportable disease in NYC, and the NYC DOH maintains a registry of all cases reported from physicians, clinics, and hospitals that meet the revised Centers for Disease Control (CDC) criteria for definite or compatible cases.[11] Additionally, the DOH maintains a birth certificate registry that contains information on all babies born in NYC and on selected characteristics of their mothers.

In the case-control study, all babies reported with congenital syphilis in 1989 were identified from the congenital syphilis registry, and the birth certificate registry was then reviewed to locate their birth certificates. All congenital syphilis babies with available birth certificates were defined as cases. The next two babies born without congenital syphilis on the same day as each case were selected as controls. Cases and controls were also matched on race and hospital of birth. If controls could not be selected on the same date of birth as a case, the subsequent day's births were reviewed until an appropriate control was located. For each of the 14 twin case sets, two controls were selected. All information collected on the babies (race, gestational age, and birth weight) and on their mothers (use of cocaine/crack, heroin, marijuana, methadone, and alcohol; level of prenatal care; age; marital status; level of education; and insurance status) was obtained from the babies' birth certificates.

A bivariate analysis was done to compare cases and controls for each baby and maternal characteristic. The strength of association was evaluated by the odds ratio with 95% confidence intervals; statistical significance was set at the $P < .05$ level. Correlations between variables were examined. Conditional logistic regression analysis was performed to assess the independent association of each of the maternal characteristics with congenital syphilis. Data were analyzed using Statistical Analysis System (SAS) version 5.18.[12]

Results

Of the 987 babies born in NYC listed on the congenital syphilis registry as of March 1990, there were 72 without available birth certificates, 32 stillbirths, and 1 with no available controls. Birth certificates were occasionally not located because of discrepancies in the legal names of the children between the birth certificate and congenital syphilis registries. Of

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The authors are with Epidemiology & Prevention Services, New York City Department of Health, New York, NY.

Requests for reprints should be sent to Mark Htoo, MD, New York City Department of Health, 125 Worth Street, PO Box 73, New York, NY 10013.

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the remaining 882 eligible babies, 881 were compatible cases of syphilis and 1 was a definite case.

Cases and controls were matched on race and included non-Hispanic blacks (73%), Hispanics (22%), Whites (4%), and others or unknown (1%). In the bivariate analysis (Table 1), cases had a significantly lower mean birth weight and gestational age than controls (both $Ps < .001$). Cases were statistically more likely to use cocaine/crack, alcohol, or other drugs; have no prenatal care; be single; and have less than a high school education (all $Ps < .001$). Cocaine/crack use was significantly correlated with alcohol use and being unmarried, and prenatal care was correlated with education (correlation coefficients of .35524, .1988, and .06817, respectively, all $Ps < .001$).

Independent of the effect of other factors, the odds of being exposed to cocaine were 3.9 times greater among cases than among controls (Table 2). Other factors significantly associated with congenital syphilis were being unmarried, alcohol use, number of prenatal visits, and having less than a high school education.

Discussion

In this investigation, a strong association was found between congenital syphilis and maternal cocaine/crack use. Moreover, the increase in congenital syphilis cases in NYC since 1985 was temporally related to an increase in cocaine/crack use (Figure 1), whereas the use of alcohol and other drugs remained relatively stable during this period.[3]

Several limitations must be considered in interpreting these data. Underreporting may be a serious problem in examining congenital syphilis in NYC.[13] Drug use information tends to be based on self-reports, and collection of information on the birth certificate is not standardized throughout hospitals. Although the possibility exists that drug histories were more actively pursued among mothers of babies with congenital syphilis, questioning for drug usage is a routine part of medical history-taking done immediately after birth and well before syphilis serologic results are available.

Maternal cocaine/crack use might contribute to the occurrence of congenital syphilis in a variety of ways. Cocaine/crack use might increase the frequency of sexual activity because of the exchange of sex for drugs, or might lead to an increase in unprotected sexual activity.

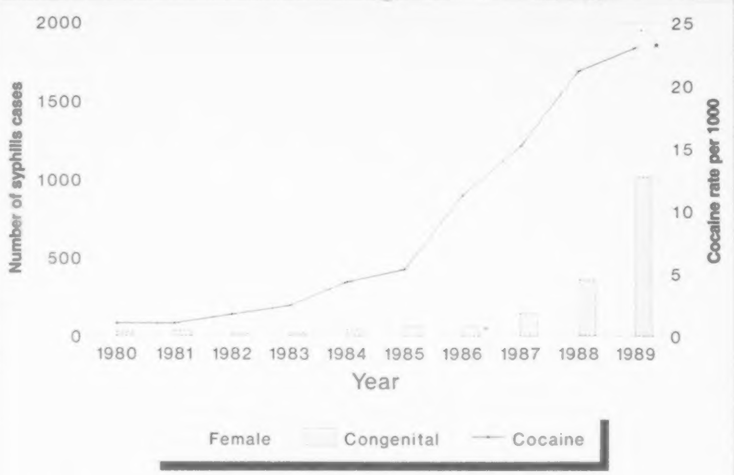

FIGURE 1—Female and congenital syphilis cases and cocaine rate per 1000 live births, New York City, 1980 to 1989

TABLE 1—Comparisons of Cases and Controls: Congenital Syphilis Cocaine/Crack Study, New York City, 1989.

	Cases[a] (n = 882)		Controls[a] (n = 1736)			
Characteristics	No./Total	%	No./Total	%	Crude Odds Ratio	95% Confidence Interval
Babies						
<2500 grams	308/882	35	211/1736	12	3.9	3.2–4.7
<37 weeks gest.	259/732	35	277/1593	17	2.6	2.1–3.2
Mothers						
Cocaine/crack use	249/865	29	93/1710	5	7.0	5.4–9.2
Alcohol use	86/865	10	36/1710	2	5.1	3.4–7.8
Any drug use	403/865	47	262/1710	15	4.8	4.0–5.8
No prenatal care	224/736	30	150/1601	9	5.3	4.1–6.8
1–5 visits	207/736	28	372/1601	23	2.0	1.6–2.5
>5 visits	305/736	41	1079/1601	67	1.0	Reference
Unmarried	748/882	85	1147/1736	66	2.9	2.3–3.6
Not high school graduate	427/834	52	629/1693	37	1.8	1.5–2.1
Age <20 years	95/881	11	251/1730	15	0.7	0.6–0.9
Medicaid	526/858	61	1050/1707	61	1.0	0.8–1.2

[a]Denominators vary because missing data were excluded.

TABLE 2—Variables Significantly Associated with Congenital Syphilis in the Conditional Logistic Regression Model: New York City, 1989

Variable	Regression Coefficient	Adjusted Odds Ratio	95% Confidence Interval
Cocaine/crack	1.354	3.9	2.8–5.3
Unmarried	0.779	2.2	1.7–2.8
Alcohol use	0.723	2.1	1.2–3.4
No. of prenatal visits	−0.098	0.9[a]	0.88–0.93
Less than high school education	0.298	1.3	1.1–1.7

[a]When entered as a continuous variable, with each decreasing prenatal visit, the odds of being a case are increased by 1.1.

Cocaine/crack use could also lead to a decreased use of prenatal care, thereby causing missed opportunities for the detection and treatment of syphilis (38% of case mothers who used cocaine/crack had no prenatal care). The lower socioeconomic status of cocaine/crack-using mothers could be associated with a decreased availability of health services for this population, or with inadequate services. Among women who gave birth to babies with congenital syphilis in NYC who were reported in 1989 and who were documented to have received prenatal care, only a small proportion were documented to have received treatment for syphilis during their pregnancy.[14] Since so few delivering women received treatment for syphilis, treatment failure is probably not an important factor in the increase of congenital syphilis cases.

The provision of accessible and adequate prenatal health care, along with education targeted to women who engage in high-risk sexual activities, is needed to confront the epidemic of congenital syphilis in NYC. ☐

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Human Papillomavirus, Gonorrhea, Syphilis, and Cervical Dysplasia in Jailed Women

Nina A. Bickell, MD, MPH, Sten H. Vermund, MD, PhD, Michelle Holmes, MD, MPH, Steven Safyer, MD, and Robert D. Burk, MD

ABSTRACT

We assessed the prevalence of human papillomavirus (HPV) by cervicovaginal lavage and Southern blot and inquired about behavioral risk factors for cervical disease and sexually transmitted diseases by interview in 114 female detainees at a large New York City jail. Of the women screened, 8% had abnormal Pap smears, 35% had HPV, 7% had gonorrhea, and 22% had serologic syphilis. Given the high rates of HPV infection and cervical cytology, Pap smears should be a routine intake procedure for incarcerated women. (*Am J Public Health.* 1991;81:1318–1320)

Introduction

Prisoners have a high prevalence of drug use, psychiatric disorders, trauma, alcohol abuse, and chronic medical conditions such as respiratory diseases and hypertension.[1,2] The number of incarcerated women in the U.S. increased 53% between 1983 and 1987.[3] At our study site the number of female prisoners has more than doubled in the last 5 years (*SS:unpublished data*).

Prevalence rates as high as 10% for cervical carcinoma in situ, 16% for syphilis, and 5% for gonorrhea have been reported in incarcerated adult women.[4–7] No studies of human papillomavirus (HPV) have been published. Because prisoners have high rates of drug use and prostitution,[3] many are likely to have been exposed to the human immunodeficiency virus (HIV) through these routes.[8] Women

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Nina Bickell, at the time of this study, was with the Montefiore-Rikers Island Health Service, NY; she currently is with the Department of Medicine, Gouverneur Hospital, NY. Sten H. Vermund, at the time of this study, was with Montefiore Medical Center, Bronx, NY; he currently is with the National Institutes of Health, Bethesda, Md. Michelle Holmes is with Cambridge Hospital, Mass. Steven Safyer is with Montefiore-Rikers Island Health Service and with Montefiore Medical Center. Robert D. Burk is with Montefiore Medical Center and with the Albert Einstein College of Medicine, Bronx, NY.

Requests for reprints should be sent to Dr. Sten Vermund, MD, PhD, Epidemiology Branch, Division of AIDS, National Institute of Allergic and Immunologic Diseases, National Institutes of Health, 6003 Executive Boulevard, Bethesda, MD 20892.

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with HIV may have higher risk of HPV-associated cervical cytologic abnormalities.[9]

We undertook a cross-sectional study of the prevalence of HPV, gonorrhea, syphilis, and cervical cytology in imprisoned women, examined selected behavioral risk factors for HPV and other sexually transmitted diseases (STDs), and looked for any association of HPV with cervical abnormalities in this population.

Methods

All female detainees at the New York City Correctional Institute for Women undergo an intake examination at the Montefiore Medical Center/Rikers Island Health Service, which provides comprehensive health care to the Rikers Island Jail complex on a contractual basis with the New York City Department of Health. During four nonsuccessive weeks in 1988, eligible female detainees were asked to participate in this study. Eligibility criteria included age less than 50 years, ability to speak English, an intact uterine cervix, history of heterosexual relations, and competency to give informed consent.

The usual medical intake evaluation includes standardized medical history and a physical including pelvic exam and laboratory tests for TB, syphilis, gonorrhea, pregnancy, and a Pap smear. Those who agreed to participate also received a structured interview before the usual intake and cervicovaginal lavage for HPV.[10] The structured interview elicited additional information about sexual and medical histories. Demographic data of participants and nonparticipants were abstracted from the standardized prison medical charts.

All lab analyses were conducted in a blinded fashion. For analytic purposes, normal Pap smears and those showing only atypia with inflammation were considered "no significant pathology"; those classified with atypia and mild dysplasia, or Class III or IV, were considered abnormal in accordance with the Bethesda cytologic classification system.[11] Syphilis serology was evaluated by rapid plasma reagent tests and, if positive, by fluorescent treponemal antibody (FTA) tests. *Neisseria gonorrhoeae* samples were obtained from the endocervix with a cotton swab and cultured on Thayer Martin plates. HPV DNA was extracted from exfoliated cervicovaginal cells collected by lavage. DNA was digested with a *Pst 1* and resultant DNA fragments were separated by gel electrophoresis. HPV DNA was detected by Southern blot analysis with a mixed probe containing radiolabelled HPV 11, 16, 18, and 31 DNA initially under low stringency and following hybridization, under moderate stringency conditions according to techniques previously published.[10] Statistical calculations included odds ratios (ORs) and Mantel-Haenszel OR_{MH} to measure strength of association.

TABLE 1—Odds Ratios (OR) of HPV Infection and Condom Use among Women with Greater Numbers of Sexual Partners in the Previous Year

Partners, y (n)	OR of Condom Use (95% CI)	OR of HPV Infection (95% CI)
0–1 (54)[a]	1.0	1.0
2–5 (34)	4.1 (0.66–26.1)	1.1 (0.4–2.9)
6–24 (14)	23.7 (6.1–92.1)	2.7 (0.8–9.6)
≥25 (11)	61.5 (13.5–280.3)	1.5 (0.4–5.6)

[a] 0–1 partners = referent group.

Note. ORs are age adjusted; CI = confidence interval; Among women with 25 or more partners in the last year, 73% used condoms and 36% were infected with HPV. Among women with 6–24 partners in the last year, 50% used condoms and 57% were infected with HPV.

Results

Of 145 detainees, 130 met inclusion criteria, and 114 (88% of eligible women) agreed to participate. The most common reason given for refusal was belief that they were not infected; the most common reason for exclusion was a history of hysterectomy. Participants' (N=114) average age was 26.6 years (standard deviation [SD] ± 6.2 years); 53% were Black, 30% Hispanic, and 17% White. Nonparticipants' (N=16) average age was 27.2 years (SD ± 6.8 years); 62% were Black, 25% Hispanic, and 13% White. Sixty-eight percent of the women acknowledged use of some drug, 53% admitted to habitual use of cocaine, and 99% smoked cigarettes.

The median number of reported sexual partners in the last year was 1 (range = 0–2500). Fifty-six percent of the women stated they knew their sexual partners had other sexual partners. The mean age at first coitus was 15 years. Forty women (35%) practiced contraception: 32 used condoms, 5 had tubal ligations, 2 used IUDs and 1 used contraceptive sponges. None used oral contraceptives. Of women using condoms, 80% used them "most" or "all the time." A trend of increased condom use was observed among women with greater numbers of sexual partners (Table 1).

STDs were common: 7% of the women (8/111, 3 missing) had positive gonorrhea cultures, and 22% (24/108, 6 missing) had positive FTA. HPV was the most prevalent STD: 35% of the participants (40/114) were infected.

Pap smears were abnormal in 9% of the women (9/104, 10 missing). There were no cases of invasive cervical cancer. Five of nine women with abnormal Pap smears (56%) were infected with HPV compared to 30 of 95 (32%) women without significant Pap smear pathology (OR = 2.7, 95% confidence interval [CI] = 0.7–10.4). HPV type was not associated with Pap smear abnormality (*R.D. Burk; unpublished data*). The common HPV type prevalence rates were 28% types 16 and 18, 10% type 42, 2% types 6 and 11, and 35% "related" types. Younger women (≤25 years) were more frequently HPV infected (OR = 2.0, 95% CI = 0.9–4.5) and were more likely than older women to have had 2 to 24 sexual partners in the previous year (OR = 2.4, 95% CI = 1.01–5.8). There was an unstable relationship between age and HPV infection when number of sexual partners was controlled for (OR_{MH} = 1.8, 95% CI = 0.8–4.1).

Discussion

In 1988, the typical female detainee at Rikers Island was a woman in her mid-20s who engaged in multiple potentially harmful behaviors, including cigarette smoking, cocaine use, sexual activity by age 15, more than one sexual partner in the last year, and low frequency of condom use. The prevalence of STDs and abnormal Pap smears among the women inmates is of a magnitude similar to that of other studies in young inner city and prison women.[12–14] The marked frequency of condom use in highly sexually active women is probably associated with prostitution. The high HPV prevalence may be related to sensitive cell sampling and DNA detection methods.[15] Although this study did not assess incidence through repeated testing, such studies would be desirable.[16]

The majority of these women did not practice safe sex; 72% had not used condoms. These rates of condom use are similar to those reported among juvenile detainees.[13] High STD rates and low use of barrier contraception are particularly disturbing given the prisoners' increased risk of infection with HIV[17,18] and the increased risk of HPV-associated cervical cytologic abnormalities in immunocompromised women.[9,19]

These data suggest the need for universal Pap smear and STD screening of women in prisons and jails, careful follow-up of women with abnormal Pap smears and those who may be HIV positive, and health education for barrier contraception and STD prevention. ☐

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Self-disclosure of HIV Infection to Sexual Partners

Gary Marks, PhD, Jean L. Richardson, DrPH, and Norma Maldonado, BA

ABSTRACT

Self-disclosure of HIV infection to current sexual partners was examined in 138 seropositive men, mostly lower income Hispanics, from Los Angeles. Forty-five percent had been sexually active since learning their serostatus. Of these, 52% had kept their infection secret from one or more partners. The likelihood of disclosure decreased in direct proportion to the number of partners. Nondisclosure occurred in conjunction with unprotected receptive and insertive anal intercourse. Potential factors promoting nondisclosure are discussed. (*Am J Public Health*. 1991; 81:1321–1323)

Introduction

Individuals who are aware they are infected with the human immunodeficiency virus (HIV) and who engage in sexual relations have a social and legal responsibility to disclose their infection to their partners.[1–3] Withholding this information increases the potential for HIV transmission because it creates a situation in which unsafe sexual activity is more likely to occur. Should a partner become infected, he or she may unknowingly infect others. This chain of events suggests that nondisclosure of HIV infection to sexual partners may be an important cause of the continuing HIV epidemic.

Little is known about the prevalence of this behavior. Studies have examined intentions to inform partners among individuals who had not been tested for HIV[4] or had not yet received test results.[5] Two studies have focused on HIV-infected individuals. Of 25 seropositive women, only 52% had disclosed their infection to all sexual partners since diagnosis.[6] One hundred and seven homosexual men informed only 31% of their new sexual partners.[7] We report findings of an ongoing study of self-disclosure in relation to the specific sexual activities performed.

Methods

Adult men who have tested seropositive for HIV in the previous 18 months were recruited for study at a public HIV clinic located in a predominantly Hispanic section of Los Angeles. Study candidates were from any ethnic group, spoke English or Spanish, and were 18 years of age or older.

Recruitment was conducted in the waiting area of the clinic by a bilingual Hispanic woman. Of the men approached, 96% agreed to participate. Each subject completed a self-administered confidential questionnaire in a private section of the waiting area. A four-digit numerical code, uniquely derived by each subject, was used as identification.

Sexual activity and disclosure were measured with a matrix checklist thoroughly explained to each subject. Instructions requested subjects to think of all partners since testing seropositive, including intimate lovers, casual partners, and spouse if married. For each partner, subjects checked the activities that occurred[8,9] and indicated whether they had revealed their HIV infection before the activity (yes = 1; no = 0) and whether the partner was known to be infected with HIV (yes = 1; no or don't know = 0). Subjects who had not had any sexual relations during this time period checked a box at the top of the checklist.

Results

The sample consisted of 138 men (104 Hispanics, 83% of whom used a Spanish language questionnaire; 21 non-Hispanic Whites; 11 Blacks; and 2 Asians). Ninety-one percent identified themselves as homosexual or bisexual. Overall, 13% reported ever injecting a nonprescription drug and 2% reported injecting drugs since they tested seropositive. Seventy percent had an annual income under $10 000.

Sixty-one percent were asymptomatic or had minimal symptoms short of AIDS-Related Complex (ARC), 16% had ARC, and 23% had AIDS.[10] Seventy-five percent had tested HIV seropositive within the previous 8 months.

Sixty-two of the 138 men (45%) had been sexually active (other than kissing) since learning their serostatus. Of these 62, 4 were deleted from the disclosure analysis due to missing data on relevant questionnaire items. Thirty out of 58 men (52%) kept their infection secret from one or more partners. Disclosure did not differ significantly by ethnicity (i.e., Hispanic vs non-Hispanic).

Of the men who had only 1 sexual partner since diagnosis (n = 36), 69% disclosed their serostatus to that partner. Among those with 2 to 4 partners (n = 11),

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The research for this paper was conducted at the Department of Preventive Medicine, University of Southern California.

Requests for reprints should be sent to Gary Marks, PhD, Department of Preventive Medicine, University of Southern California, Parkview Medical Building, A-301, 1420 San Pablo Street, Los Angeles, CA 90033.

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TABLE 1—Sexual Activity since Testing HIV Seropositive and Disclosure of HIV Infection to Partners

	Nondisclosure				Disclosure			
	All Partners		HIV-Neg. Partners		All Partners		HIV-Neg. Partners	
Sexual Activity	Ss	Ps	Ss	Ps	Ss	Ps	Ss	Ps
Anal insertive/NC	10	13	10	13	12	13	5	5
Anal insertive/C	12	23	12	22	12	15	9	12
Anal receptive/NC	17	41	17	41	8	8	3	3
Anal receptive/C	16	51	16	48	13	15	10	12
Oral insertive/NC	17	40	14	35	12	15	6	8
Oral insertive/C	5	9	5	9	5	5	4	4
Oral receptive/NC	21	71	21	69	14	15	8	8
Oral receptive/C	7	26	7	26	4	4	4	4
Vaginal insertive/NC	3	3	2	2	0	0	0	0
Vaginal insertive/C	5	15	5	15	4	4	4	4

Note: Ss = no. of subjects; Ps = no. of partners; NC = without condom; C = with condom. Sixty-two men reported sexual activity other than kissing since testing seropositive. Four men from this subgroup had missing data on disclosure and partners' HIV status and were deleted from the analysis. Thus, the entries can be evaluated against 58 men who had sexual relations since testing seropositive and provided full data or against the total sample of 138 men.

36% disclosed to at least 1 partner. Among those with 5 or more partners (n = 11), only 18% disclosed to at least 1 partner. This highly significant trend ($(\chi^2) = 9.50$, $P < .01$) indicates that the likelihood of disclosure decreased in direct proportion to the number of partners.

Table 1 displays, for each sexual activity, the total number of subjects, total number of partners, and number of partners presumed to be seronegative at the time of the activity. Information is provided separately for partners to whom subjects disclosed and those to whom subjects did not disclose. Ten subjects (17% of the sexually active) engaged in insertive anal intercourse without using a condom and without disclosing to 13 seronegative partners. Seventeen men (29% of the sexually active) engaged in unprotected receptive anal intercourse without disclosing to 41 seronegative partners. The higher prevalence of receptive than insertive anal intercourse is consistent with a study of HIV-infected homosexual White men from Boston.[11] Both activities are risk factors for transmitting the virus.[12]

In Table 1, the difference between the total number of partners and the number of seronegative partners represents the number of seropositive partners. This difference was larger for informed than uninformed partners, indicating that subjects tended to reveal their infection to those known to be seropositive. When the activity involved insertive or receptive anal intercourse, disclosure to seropositive partners generally occurred in combination with unprotected contact, whereas disclosure to seronegative partners generally occurred in combination with protected contact.

Finally, we examined whether the gender of the partner was associated with disclosure among the bisexual men. Because the activity checklist did not explicitly ask subjects to indicate partner gender, we compared nondisclosure rates for vaginal intercourse (female partners) with nondisclosure rates for receptive anal intercourse (male partners). No significant differences were found.

Discussion

In a sample consisting primarily of lower socioeconomic homosexual/bisexual Hispanic men, we found a relatively high prevalence of nondisclosure of HIV infection to current sexual partners in combination with unsafe sexual activity. For this sample, nondisclosure was not due entirely to lack of knowledge about sexual transmission of HIV. Ninety-four percent of the subjects were aware that they could infect others through unprotected sexual intercourse. Use of alcohol or other nonintravenous drugs before sexual relations, as well as attitudes about social responsibility, may have been factors, although we do not have confirmatory information.

The restricted nature of our sample precludes any wide generalization of the results. Additional studies of Whites, Blacks, Hispanics, and other minority groups of varying socioeconomic levels are needed before a more refined picture can be obtained. Investigation of factors promoting nondisclosure would be immensely valuable for counseling and educational purposes. ☐

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Immunoglobulins and Growth Parameters at Birth of Infants Born to HIV Seropositive and Seronegative Women

Arlene Butz, CPNP, ScD, Nancy Hutton, MD, and Elaine Larson, RN, PhD

ABSTRACT

Health characteristics of neonates born to human immunodeficiency virus (HIV) seropositive (n=63) and HIV seronegative (n=57) women were compared. No significant differences were found between the two infant groups for sociodemographic characteristics, mode of delivery, Apgar scores, gestational age, growth parameters, or immunoglobulin levels. Furthermore, no differences were detected in a subgroup of infants later confirmed to have HIV infection (class P-2). Maternal HIV infection does not appear to affect newborn health characteristics. (*Am J Public Health.* 1991;81:1323–1326)

Introduction

As of July 1990, 2464 cases of acquired immunodeficiency syndrome (AIDS) had been identified in patients under 13 years of age,[1] with 78% to 80% presumed to have contracted human immunodeficiency (HIV) infection by vertical transmission.[2,3] The efficiency of virus transmission from an HIV-infected mother to infant is reported to be 13% to 50%.[4–6] Early diagnosis of infection is difficult because transplacentally transmitted anti-HIV antibodies are present in all infants of HIV-seropositive mothers and persist up to 15 months of age or longer.[4,7]

Elevated serum immunoglobulin (IgG, IgM, and IgA) levels are a hallmark of pediatric HIV infection, often developing prior to other laboratory abnormalities such as a decrease in T helper (CD4) cells.[8,9] Birth weight, head circumference, and gestational age have been inconsistently reported to be decreased in infants with perinatal HIV infection.[10–13]

This ongoing prospective study followed a cohort of infants born to mothers at risk for HIV infection to compare the health and growth characteristics of HIV-infected and HIV-noninfected infants from comparable risk strata. The analysis of birth characteristics, specifically growth parameters and immunoglobulins, is presented here.

Methods

All infants born to HIV-seropositive women delivering at an inner-city hospital were eligible for study. An equal number of HIV-seronegative women with acknowledged risk behaviors (intravenous [IV] drug use, sexual contact with IV drug user) were enrolled as control subjects. HIV counseling and testing are offered to all obstetric patients who deliver at this hospital. For most mothers, maternal risk status was determined prenatally by standardized questionnaire and HIV testing performed for consenting women. Newborn HIV testing was offered to unregistered mothers who acknowledged HIV risk factors. Overall there is a 5% seropositivity and 10% refusal rate in this population.[14] In order to examine differences in appropriate intrauterine growth independent of prematurity, infants with less than 34 weeks gestation were excluded.

Each infant's blood was tested within 2 weeks of birth (usually within 24 hours of birth) for HIV antibody and quantitative immunoglobulins, and medical records were reviewed. Weight, length, head circumference, and gestational age were measured in a standardized fashion immediately postdelivery by designated nursery staff. Weight and gestational age by Ballard score[15] was used to classify newborns for intrauterine growth:[16] small, appropriate, or large for gestational age (SGA, AGA, LGA, respectively).

IgG and IgM quantitation was performed by nephelometry (a light scattering technique). Radioimmunodiffusion was used to quantitate IgA levels. The presence of HIV antibody was determined by ELISA (Genetics System, Seattle, Wash) and confirmed by Western blot (DuPont, Wilmington, Del).

Results

Between August 4, 1988, and June 1, 1990, 120 of 127 mothers (95%) agreed to enroll their infant in the study, including 10 infants born to unregistered mothers (5 HIV positive and 5 HIV negative). Sixty-three infants (52.5%) were born to HIV-seropositive mothers, 91% were Black, and 49% were male (Table 1). Maternal substance abuse was the predominant

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Arlene Butz and Elaine Larson are with the School of Nursing, and Nancy Hutton is with the Department of Pediatrics, School of Medicine, of The Johns Hopkins University.

Requests for reprints should be sent to Arlene Butz, CPNP, ScD, The Johns Hopkins University/School of Nursing, 600 N Wolfe Street, Baltimore, MD 21205.

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TABLE 1—Sociodemographic and Health Characteristics of Infants by HIV Serostatus at Birth (n=120)[a]

Characteristics	HIV Positive (n=63)	HIV negative (n=57)	Total No. (%)
Sociodemographic			
Maternal age (years)	26.2	25.9	26.1
Race			
Black	60	50	110 (91.7)
White or other	3	7	10 (8.3)
Sex of infant			
Male	35	24	59 (49.2)
Female	28	33	61 (50.8)
Risk factor			
IVDA[b]	44	46	90 (75.0)
Sex partner	14	6	20 (16.6)
Multi-risk or unknown	5	5	10 (8.4)
Prenatal Substance Abuse			
Yes	42	48	90 (75.0)*
Health			
Mode of Delivery			
Vaginal	54	47	101 (84.2)
C-section	9	9	18 (15.0)
Missing data			1 (.8)
Apgar score (mean)			
1 minute	7.96	7.98	
5 minute	8.76	8.67	
Gestational age (mean weeks)	38.36	38.66	

[a] Differences between groups for all parameters were nonsignificant at $p \geq .14$ by *t* test and χ^2 test.
[b] IVDA=intravenous drug abuser.
**P*=.06.

TABLE 2—Mean Growth Parameters at Birth by Infant HIV Serostatus (n=120)

Parameter	HIV Positive (n=63)	HIV Negative (n=57)	*t* test	*P* Value
Weight (g)	2811	2878	−.67	.51
Length (cm)	47.91	47.94	−.05	.96
Head circumference (cm)	33.23	32.77	1.17	.24
(range)	(28.5–43.5)	(28–37)		

TABLE 3—Mean (range) Immunoglobulin Levels (mg/dl) at Birth by HIV Serostatus

Immunoglobulin	HIV Positive	HIV Negative	*t* test	*P* Value
IgG	1385 (494–2190)	1305 (370–1760)	1.29	.20
IgM	16.4 (6–47)	13.5 (6–59)	1.36	.18
IgA	3.92	3.76	.28	.78
	Normal Values			
	Meites		Fosarelli[b]	
IgG	1086 ± 290		349–1701	
IgM	14 ± 6		16–72	
IgA	2 ± 2		0–13	

[a] Data from Meites S.[16]
[b] Data from Fosarelli et al.[17]

HIV risk factor (75%). Prenatal drug exposure occurred in 84% of infants born to HIV-negative women vs 67% of infants born to HIV-positive women (*P*=.06). Type of delivery (85% vaginal) was not statistically associated with seropositivity. Prenatal sexually transmitted disease (gonorrhea and syphilis) exposure was not significant between HIV-positive and HIV-negative infants (*P*=.24). There was no difference between groups for mean Apgar scores at 1 and 5 minutes or mean gestational age (Table 1).

Fourteen infants (six HIV-positive) were SGA. The majority of these infants (79%) were born to prenatally substance-abusing mothers (HIV-negative =7/8, HIV-positive =4/6; Fisher's exact, *P*=.39). For all infants, there was no significant difference between SGA and non-SGA infants for prenatal drug use (Fisher's exact, *P*=.54).

There was no statistically significant difference for any growth parameter between HIV-positive and HIV-negative infants (Table 2). Mean birth weight for HIV-positive infants was 2811 g and 2878 g for HIV-negative infants (range for all infants: 1585–4586 g, *P*=.51). Three infants (one HIV-positive, two HIV-negative) had a birth weight between 1500 g and 1800 g.

Mean differences for quantitative IgG, IgM, and IgA between the HIV-positive and HIV-negative infants were not statistically significant (Table 3). When mean values for IgG, IgM, and IgA in each group of infants were compared to standardized values previously published[17,18] for a comparable population of Black infants from the same institution,[16] all were within normal limits except for the mean IgM level for HIV-negative infants (13.5 mg/dl), which was slightly lower.

Ten HIV-positive infants developed class P-2 HIV disease[19] during the study period (range of follow-up=9–30 months). Comparison of birth weight, gestational age, and immunoglobulin levels of these 10 HIV-infected infants with remaining HIV-positive infants (P-O, i.e., indeterminate status) and HIV seroreverters (infants born HIV antibody seropositive who reverted to seronegative status by 15–18 months of age) revealed no differences among the three groups (Table 4) with the exception of higher IgA in the HIV reverter group.

Discussion

Our data indicate that in a US urban center with a high prevalence of IV drug use, birth characteristics did not differ between infants born to HIV-positive mothers and HIV-negative mothers with comparable risk. Immunoglobulins were not significantly different between the two groups of infants or between HIV indeterminate, HIV-infected, or HIV seroreverter infants. Elevated immunoglobulin

TABLE 4—Selected Health Characteristics at Birth by HIV Infection Classification[a]

Characteristic	P-2[b] (n=10)	P-O[c] (n=39)	Reverters[d] (n=14)	HIV negative[e] (n=57)	F Test (*P* value)	Difference (*P* value)
Growth						
Birth weight (g)	3049 (2724–3374)	2778 (2612–2945)	2721 (2436–3006)	2878 (1770–4586)	1.35 (.27)	NS
Head circumference (cm)	33.2 (31.7–34.8)	33.2 (32.4–33.9)	33.4 (31.9–34.9)	32.77 (28–37)	.03 (.97)	NS
Gestational age (weeks)	38.9 (37.7–40.1)	38.2 (37.5–38.8)	38.4 (37.3–39.5)	38.7 37.6–38.9)	.52 (.59)	NS
Immunoglobulins (mg/dl)						
IgG	1303 (1049–1556)	1430 (1301–1559)	1324 (1122–1527)	1305 (370–1760)	.65 (.53)	NS
IgM	17.4 (9.4–25.5)	14.04 (9.9–18.1)	21.7 (15.3–28.1)	13.5 (6–59)	2.11 (.13)	NS
IgA	3.3 (.9–5.7)	3.0 (1.9–[illegible].1)	6.6 (4.8–8.4)	3.76 (0–9)	5.91 (.006)	p=.05

[a] Numbers are means and, in parentheses, 95% confidence intervals.
[b] P-2=symptomatic HIV infection.
[c] P-O=indeterminate status (HIV antibody seropositive at birth, asymptomatic <15 months of age).
[d] Reverters=HIV antibody seropositive at birth and converted to HIV seronegative by 15–18 months of age.
[e] HIV negative=HIV seronegative at birth and remained seronegative (controls).

levels have been found by 5 to 9 months of age in HIV-infected infants.[9,20] We found no retardation in growth that could be attributed to maternal HIV infection when infants with similar environmental exposures were compared. Risk factors for HIV infection were dominated by maternal IV drug use or sexual contact with an IV drug user, comparable to trends for HIV infection in women in Western countries.[21–23] Because screening programs targeted to high-risk women fail to identify a substantial proportion of women with HIV infection[14,24] and prenatal HIV testing is voluntary, it is possible that some HIV-exposed infants were not included in this sample. The sample size achieved was sufficient to detect clinically significant differences in weight (265 g), length (1.43 cm), and head circumference (1.38 cm) between HIV-positive and HIV-negative groups with a power of 80%.[25]

Neither set of birth characteristics, growth parameters, or immunoglobulins distinguishes HIV-infected infants from comparable risk noninfected infants at birth. All infants born to HIV-risk women require vigilant pediatric follow-up for early diagnosis of true HIV infection. ☐

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The Impact of HIV on the Usefulness of Sputum Smears for the Diagnosis of Tuberculosis

Richard Long, MD, Marcella Scalcini, MD, MPH, Jure Manfreda, MD, Michel Jean-Baptiste, MD, and Earl Hershfield, MD

ABSTRACT

In a developing country, 289 patients were examined for active pulmonary mycobacterial disease (sputum smear and culture) and HIV infection (serology) to compare the sensitivity and positive predictive value of sputum smears for diagnosing pulmonary tuberculosis in patients with and without antibodies to HIV. Seventy-nine percent of HIV-seronegative vs 66% of HIV-seropositive patients with positive cultures for *Mycobacterium tuberculosis* were smear positive ($P < .05$), and a positive sputum smear predicted the presence of *M. tuberculosis* in 90% of HIV seronegative vs 80% of HIV seropositive patients ($P < .05$). In our opinion, HIV did not significantly compromise the diagnostic utility of the sputum smear. (*Am J Public Health.* 1991;81:1326–1328)

Introduction

The continuing usefulness of sputum smears for the diagnosis of tuberculosis in developing countries has been questioned. In a recent editorial,[1] Pitchenik expressed concern that immunosuppression resulting from human immunodeficiency virus (HIV) type 1 may not only have reduced the sensitivity of the sputum smear by reducing caseation necrosis, and thus the number of acid-fast bacilli in the airway,[2-5] but may also have affected the specificity of the sputum smear by increasing the proportion of patients with nontuberculous mycobacteria.[6] To determine whether HIV has compromised the usefulness of the sputum smear, we compared the sensitivity and positive predictive value of smears for the diagnosis of pulmonary tuberculosis in rural Haitians with and without antibodies to HIV.

Methods

Over a 1-year period beginning March 1988, 323 consecutive rural Haitians presented to a district hospital (Hopital Albert Schweitzer) with suspect pulmonary mycobacterial disease had positive sputum smears for acid-fast bacilli and/or positive cultures for mycobacteria. None had received antituberculous drugs in the preceding 12 months. HIV antibody status was determined in 297 (92%) patients; serum was not available for 4 patients who died shortly after admission and 22 who either refused the test or had already left the facility. Eight HIV-tested children were excluded because the positive culture was from a gastric aspirate and not a sputum specimen. This left a total of 289 (89%) patients for whom results of HIV testing, sputum smear examination, and sputum culture were available.

Two to three fresh, first morning deep cough specimens of sputum were collected from each patient and processed in a manner identical to that reported previously.[7] Standard methods of decontamination, staining, culture, and speciation were applied.[8-11] After obtaining informed consent, serum was collected and tested for antibody to HIV using standard methodology.[12]

Chi-square and Fisher's Exact test were used to compare the sensitivity and positive predictive value of sputum smears for the diagnosis of tuberculosis in HIV-seropositive and HIV-seronegative patients with pulmonary mycobacterial disease.[13]

Results

The 289 study patients ranged in age from 12 to 95 years; 54% were male and 46% female. Table 1 shows that 74 (26%) were HIV seropositive, 222 (77%) smear positive, and 266 (92%) culture positive. Culture was positive for *Mycobacterium tuberculosis* alone in 236, for both *M. tu-*

The research for this paper was conducted at the Department of Medicine, Hopital Albert Schweitzer, Haiti (MS, M J-B), and the Department of Medicine, University of Manitoba, Winnipeg, Canada.

Requests for reprints should be sent to Richard Long, MD, Room C5116, Department of Medicine, St. Boniface General Hospital, 409 Tache Avenue, Winnipeg, Manitoba, Canada, R2H 2A6.

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berculosis and nontuberculous mycobacteria (NTM) in 21, and for NTM alone in 9 patients.

It can be deduced from Table 1, that the sputum smear was significantly ($P<.05$) less likely to be positive for acid-fast bacilli in HIV seropositive (50/74, 68%) than in HIV seronegative (172/215, 80%) patients. Likewise, a sputum culture positive for *M. tuberculosis* was less likely ($P=.05$) to be present in HIV seropositive (61/74, 82%) compared to HIV seronegative patients (196/215, 91%). In contrast, HIV seropositive patients were more likely ($P=.05$) to have sputum cultures positive for an NTM alone (5/74, 7%) than were HIV seronegative patients (4/215, 2%). The proportion with a negative culture was not significantly different between HIV seropositive (8/74, 11%) and seronegative (15/215, 7%) patients.

TABLE 1—Sputum Smear and Culture Status of HIV Seropositive and Seronegative Haitians with Pulmonary Mycobacterial Disease

	HIV Positive (n=74)		HIV Negative (n=215)	
Culture	Smear +	Smear −	Smear +	Smear −
TB	37	18	143	38
TB and NTM	3	3	12	3
NTM	2	3	2	2
Negative	8	0	15	0
Total	50	24	172	43

Note. HIV = human immunodeficiency virus type 1; TB = *M. tuberculosis*; NTM = nontuberculous mycobacteria; + = positive; − = negative.

TABLE 2—Species of NTM Isolated from HIV Seropositive and Seronegative Haitians with Pulmonary Mycobacterial Disease

	HIV Positive				HIV Negative				
	Smear +		Smear −		Smear +		Smear −		
	TB Culture		TB Culture		TB Culture		TB Culture		
NTM	+	−	+	−	+	−	+	−	Total
M. scrofulaceum	2	2	1	1	6		1		13
M. gordonae					4		2		6
M. fortuitum				1[a]		2[b]		2[a]	5
M. flavescens			1		1				2
M. avium complex					1				1
M. vaccae			1						1
M. avium complex M. gordonae	1								1
M. gordonae and *M. fortuitum*				1[a]					1
Total	3	2	3	3	12	2	3	2	30

Note. HIV = human immunodeficiency virus type 1; NTM = nontuberculous mycobacteria; TB = *M. tuberculosis*; + = positive; − = negative.

[a] Biovariant *peregrinum*.

[b] Biovariant *fortuitum*.

The sensitivity of a positive acid-fast smear to identify subjects with a positive *M. tuberculosis* culture was significantly greater in HIV seronegatives than in HIV seropositives. The sensitivity in HIV seronegatives was 79% (155/196) in comparison to 66% (40/61) in HIV seropositives. In addition, a positive sputum smear predicted the presence of *M. tuberculosis* in 80% (40/50) of HIV seropositive patients and 90% (155/172) of HIV seronegative patients, a difference that was also statistically significant ($P<.05$). Neither the sensitivity nor the positive predictive value varied by age or sex.

The species of NTM isolated from 30 patients are shown in Table 2. *M. scrofulaceum* was the most common NTM isolated (13/30, 43%). In 10 instances it was isolated together with *M. tuberculosis* and in 3 it was the only isolate. All of the latter isolates were from HIV seropositive patients. *M. fortuitum* was the next most common NTM isolated (6/30, 20%). In each instance (except one where it was isolated together with *M. gordonae*) it was the only isolate, accounting for 67% (6/9) of the patients in which an NTM alone was isolated. Of the 6 patients from whom *M. fortuitum* was isolated, 2 were HIV seropositive.

Discussion

Particularly in developing countries, where resources for mycobacterial culture are often not available, acid-fast smears are relied on to diagnose tuberculosis. Before the AIDS epidemic the sensitivity of a positive sputum smear was relatively low, but the positive predictive value was high.[14] A patient with a positive smear was almost certain to have tuberculosis and was treated accordingly. Whether HIV modifies the sensitivity and positive predictive value of the sputum smear, especially in developing countries where HIV seroprevalence may be high,[15] is not known. In our study of rural Haitians we found small but significant differences in the sensitivity and positive predictive value of sputum smears between HIV seropositive and HIV seronegative pulmonary tuberculosis patients; 79% of HIV seronegative vs 66% of HIV seropositive patients with positive cultures for *M. tuberculosis* were smear positive, and a positive sputum smear predicted the presence of *M. tuberculosis* in 90% of HIV seronegative vs 80% of HIV seropositive patients. In our opinion, these differences are small and do not significantly compromise the usefulness of sputum smears in developing countries such as Haiti, where the prevalence of tuberculous infection is high.[16] Regardless of HIV status, patients with sputum smear positive pulmonary mycobacterial disease should be assumed to have disease caused by *M. tuberculosis* and should be treated accordingly. ☐

Acknowledgments

The authors thank George Carre, MD, and Roosevelt Occenac of Hopital Albert Schweitzer; Shirley Parker, MD, Joyce Wolfe, and Brenda Binda of the Mycobacteriology Reference Laboratory Health Sciences Centre, Winnipeg, Manitoba; Albert Laszlo, PhD, and Dorothy Helbecque of the Laboratory Centre for Disease Control, Ottawa, Ontario, Canada; and Eileen Walters and Heather Smith for the preparation of the manuscript. This work is dedicated to the memory of William Larimer Mellon, Jr. (1910–1989), who together with his wife Gwen, founded Hopital Albert Schweitzer in Haiti.

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Human Exposure to Rabies from Pet Wild Raccoons in South Carolina and West Virginia, 1987 through 1988

Bradley A. Woodruff, MD, MPH, Jeffrey L. Jones, MD, MPH, and Thomas R. Eng, VMD, MPH

ABSTRACT

During 1987 and 1988, exposures to eight pet wild raccoons in South Carolina and West Virginia resulted in administration of rabies postexposure prophylaxis to 19 children and 26 adults. All eight raccoons appeared normal at the time of capture, and three had no signs of illness when sacrificed. The direct medical cost resulting from these exposures was $23 714 ($527 per person). Regulations and public education may help decrease this type of rabies exposure. (*Am J Public Health.* 1991;81:1328–1330)

Introduction

Although human rabies deaths in the United States are extremely rare, approximately 18 000 persons per year receive rabies postexposure prophylaxis (CDC, unpublished data). In most southeastern and mid-Atlantic states, including South Carolina and West Virginia, raccoons account for the largest proportion of laboratory-confirmed rabid animals.[1] Most human exposures from rabid raccoons and other wild animals involve animals encountered in the wild, but exposures to rabid wild animals kept as pets have been documented.[2–4] As a result, several health organizations and government agencies recommend that wild animals not be kept as pets, and most states have laws requiring special permits to keep such animals[5]; however these measures may be ineffective. In this report, we describe eight incidents where wild raccoons were captured and kept as pets resulting in potential human exposure to rabies in South Carolina and West Virginia in 1987 and 1988.

Methods

State public health personnel in West Virginia and South Carolina were interviewed regarding incidents of possible human rabies exposure resulting from pet wild raccoons during 1987 and 1988. All reports were further investigated by interviewing exposed persons, physicians, and public health personnel and by reviewing records in state laboratories, physicians' offices, and emergency rooms. Data were available only for persons who received prophylaxis.

Bradley A. Woodruff and Jeffrey L. Jones are with the Epidemiology Program Office of the Centers for Disease Control. Bradley A. Woodruff also is with the West Virginia Department of Health, and Jeffrey L. Jones also is with the South Carolina Department of Health and Environmental Control. Thomas R. Eng is with the Division of Viral and Rickettsial Diseases of the Center for Infectious Diseases.

Requests for reprints should be sent to Bradley A. Woodruff, NCID:DVRD:HB, Mailstop A-33, Centers for Disease Control, 1600 Clifton Road, Atlanta, GA 30333.

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Human exposure to rabies was defined as a bite, a scratch, or direct contact with saliva from a laboratory-proven rabid animal. Potential exposure was defined as similar contact with an animal that was not tested.

Direct medical care costs included the cost of rabies immune globulin (RIG) and human diploid cell vaccine (HDCV) and the costs for services provided by hospitals and physicians. Medical costs were calculated from billed amounts. Direct social costs, such as the costs of contact investigations and laboratory testing, and indirect personal costs, such as lost wages and transportation costs, were not included in cost estimates because data were unavailable.

TABLE 1—Summary Data for Eight Incidents of Actual and Potential Human Rabies Exposures from Pet Wild Raccoons, South Carolina (SC) and West Virginia (WV), 1987–1988

	Number of Persons Exposed		Type of Exposure			
State	Child[a]	Adult	Bite	Scratch or Contact with Saliva	Duration of Captivity (months)	Minimum Incubation Period[b] (months)
SC	5	8	7	6	1	1
SC	0	2	1	1	.25	.25
SC	1	0	1	0	?[c]	?[c]
SC	2	2	4	0	1	1
SC	1	1	2	0	12	1.25
WV	7	9	1	14	1	1
WV	0	1	1	0	8	3
WV	3	3	1	4	0.5	0.5
Total	19	26	18	25	3.4 (mean)	1.1 (mean)

[a] ≤15 years of age.

[b] Time between last possible contact with a potentially rabid animal and time of the raccoon's diagnosis of rabies (either onset of signs or date of laboratory testing).

[c] Unknown.

Results

During 1987 and 1988, 8 incidents occurred in South Carolina and West Virginia involving potential or actual human rabies exposure from pet wild raccoons (Table 1). None of the eight raccoons had signs of rabies or behaved abnormally at the time of capture. The median interval between capture and the discovery of rabies was 1 month (range = 7 days to 1 year). The precise incubation period was unknown for these animals; however, minimum incubation periods could be derived from the date of last possible contact with other potentially rabid animals.

Six raccoons were tested for rabies; five were positive and one was reported as "inconclusive." Of the five positive animals, two appeared normal at the time of sacrifice, two died after developing signs of illness, and one died suddenly without signs of illness. Two animals were not tested for rabies: one was shot in the head, making testing impossible, and the other escaped after biting the owner. Of the three animals sacrificed without signs of illness, two were killed only after persons who had been bitten demanded that the animals be tested.

Of the 45 persons receiving prophylaxis, 19 (42%) were children ≤ 15 years of age. Eighteen (40%) persons were bitten, 25 (56%) were scratched or exposed to saliva, and 2 (4%) had no apparent exposure but demanded prophylaxis. All were residents of rural areas.

Five raccoons exposed only members or close friends of a single household. Three raccoons exposed others in the community. In the first such incident, a baby raccoon was passed between two families and handled by many neighborhood children, exposing 16 persons. In the second incident, a rabid raccoon displayed in a store exposed 13 customers and staff. In the third incident, a science teacher encouraged a student to bring his pet raccoon to school; two students in the class required prophylaxis after the raccoon was found to be rabid.

The 45 persons in the 8 incidents received 420 cc of RIG and 225 doses of HDCV at a cost of $21 092. Physician, emergency department, and clinic fees were $2622, for total direct health care costs of $23 714 or $527 per treated person.

Discussion

Keeping wild animals as pets poses a substantial risk of human rabies exposure. No vaccine can prevent rabies infection in such animals, because current rabies vaccines are not licensed for use in wild animals. As in all the incidents described in this report, owners of pet rabid animals are often not aware of their animal's infection because, at the time of capture, the animal may appear completely healthy. In addition, rabies infection in raccoons and other wild animals may have a longer subclinical course than in domestic animals.[6-8] In our series, five of eight animals had incubation periods lasting 1 month or longer. If an animal is infectious for a long period of time, both the number of persons contacting the animal and the likelihood of rabies exposure increase.

On average, a rabid pet wild animal exposes more persons to rabies than a rabid domestic animal or free-living wild animal.[9] This is probably due to longer human contact with pet wild animals than with wild animals in nature, and/or the greater attraction posed by an unusual pet such as a raccoon than is posed by more common domestic animals. In addition, pet wild animals may bite or scratch more than domestic animals, and such behavior may be more readily considered normal, resulting in delayed testing for rabies. As an extreme example of the large numbers of persons who can be exposed to a single animal, one pet rabid raccoon in Florida accounted for the prophylaxis of 172 persons at a cost of more than $64 000.[4] Moreover, many rabies exposures from wild animals may go unrecognized and untreated in children. Most dog bite injuries in this age group are not reported to health professionals[10]; the same may be true of bite injuries from wild animals.

Current legislation or regulation may be insufficient. For example, South Carolina regulations do not forbid keeping wild animals as pets. While West Virginia law allows wild animal adoption only during the hunting season for that particular species and requires a special permit, the owners of the three raccoons in West Virginia reported ignorance of these regulations. Public education is needed regarding the dangers of keeping wild animals as pets and the existence of applicable laws and regulations. Programs should target groups at highest risk, including children and residents of rural areas. ☐

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Routine Nutrition Screening Could Prevent Serious Health Problems in the Nation's Elderly, Report Says

A significant percentage of older Americans are at risk of malnutrition which could be detected and prevented through routine, inexpensive nutrition screening and greater attention to nutritional needs, according to a research report released recently.

"This excellent report lends new weight to the proposition that routine nutrition screening and care can help prevent serious illness and keep older Americans healthy and independent," said Representative Edward Roybal (D-Ca), Chair of the House Select Committee on Aging, who joined in releasing the report. "That's not only good for individuals, it's good news for a nation struggling to hold down health care costs."

The report, *Screening Older American's Nutritional Health: Current Practices and Future Possibilities,* was written by Johanna T. Dwyer, DSc, RD, Director of the USDA Frances Stern Nutrition Center at Tufts University in Boston. It is one of the most comprehensive reviews ever undertaken to examine the prevalence of malnutrition and nutrition-related problems among this nation's elderly.

"Using widely accepted criteria, a substantial proportion of older Americans have dietary intakes or diseases which place them at high risk of malnutrition," Dwyer said. "Routine nutrition screening, not only to detect the presence of malnutrition but to identify those at risk of poor nutrition, is the essential first step to improving people's nutritional status and keeping older Americans free living and in good health."

In a Foreword to the Report, Joyce T. Berry, PhD, US Commissioner on Aging, says, "Nutrition is one of many important factors affecting the health and longevity of older persons. For many low-income and minority older Americans it might be the most important."

Representative Ralph Regula (R-OH), a senior minority member of the House Select Committee on Aging, applauded Dwyer's recommendation that "screening for risk factors of malnutrition in older Americans does not need to be—nor should it be—limited to health professionals.

"Dr. Dwyer makes the case convincingly that attention to nutritional status needs to become as routine as taking our temperature or pulse," Regula said.

"The fact that nonmedical professionals and individuals themselves can employ screening techniques represents tremendous potential to keep people healthy without resorting to costly medical treatments," Regula said. "And the cost for a few minutes spent screening in a physician's office as part of a routine check-up could save many times that much in acute or chronic care treatment later."

Research has shown that people suffering from poor nutritional status have longer and more expensive hospital stays, and suffer more major complications when they are ill or injured.

Dwyer's research was commissioned by the Nutrition Screening Initiative, a five-year project focused on nutrition screening and better nutritional care among the nation's older adults.

Canine Seroprevalence and the Distribution of *Ixodes dammini* in an Area of Emerging Lyme Disease

Peter W. Rand, MD, Robert P. Smith, Jr, MD, MPH, and Eleanor H. Lacombe, BS, MT(ASCP)

ABSTRACT

This study evaluates the relative usefulness of canine serosurveys to predict risk of exposure in an area of emerging Lyme disease by comparing the distribution of canine seroprevalence with that of vector ticks. From 16 veterinary clinics throughout the State of Maine, 828 canine sera were obtained during the heartworm-testing months of April and May 1989 and measured for anti-*Borrelia* antibodies by enzyme-linked immunosorbent assay. In the same year, 1605 ticks, including 585 *Ixodes dammini*, were collected from pets, humans, small mammals, and deer. Thirty-six dogs were seropositive, 28 of which had not traveled to endemic areas. Eighty-nine percent of all seropositive dogs were from towns within 20 miles (32 km) of the coast; the great majority lived within 5 miles (8 km) of tidewater (odds ratio = 4.45, $P = .002$). Positivity varied from 17% in a southern coastal clinic to 0% in four northern clinics. Of 585 *I. dammini* identified, all but 5 (99.1%) were also from towns within 20 miles of the coast. Comparison of *I. dammini* submissions with those of another commonly found tick, *Ixodes cookei*, corroborated this predominantly coastal distribution. Canine seropositivity generally coincided with this coastal range. These data predicted areas of risk for human Lyme disease, although the prevalence of reported cases remained low. (*Am J Public Health.* 1991;81:1331–1334)

Introduction

Predicting potential areas of Lyme disease risk poses a difficult challenge for public health personnel, particularly in emerging areas with low case prevalence. Passive surveillance for reported human cases is likely to be both insensitive and of questionable specificity.[1–4] Tick vector distribution surveys based on small mammal trapping, flagging, or examination of deer are labor intensive and do not provide relative estimates of the risk of Lyme disease transmission. On the other hand, given that dogs are at greater risk of exposure than humans,[5] canine anti-*Borrelia* serosurveys offer a promising means for targeting areas for potential human risk.

The tick vector of Lyme disease, *Ixodes dammini*, was first reported in Maine in 1985.[6] Despite its subsequent establishment in coastal areas and the proximity of a hyperendemic community to the south,[7] only six cases of locally acquired Lyme disease had been reported through 1988 (unpublished data, Maine Bureau of Health Epidemiology Program, February 1991). In that year we initiated tick distribution surveys and a canine serosurvey as methods of projecting local risk of infection that might be more sensitive, efficient, timely, and economical than those involving human surveillance.

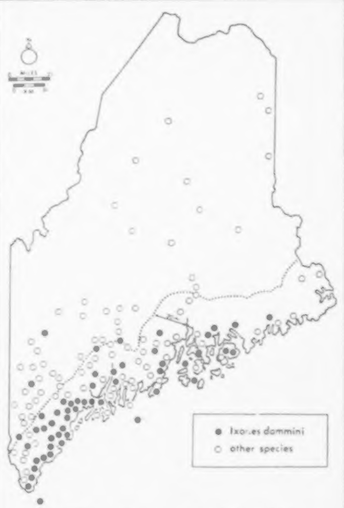

FIGURE 2—Towns of origin of identified ticks.

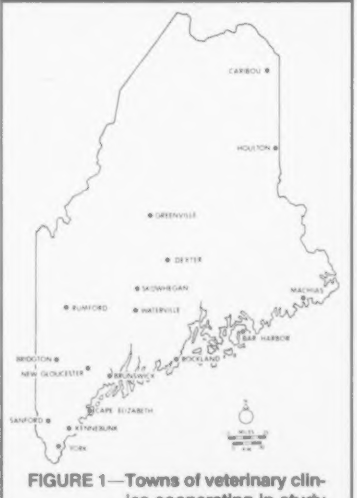

FIGURE 1—Towns of veterinary clinics cooperating in study (Dexter clinic also serves Greenville).

Materials and Methods

Ticks were submitted by veterinarians, physicians, and the public in response to published solicitations. They were also collected from natural hosts (live-trapped mice and tagged deer) by field workers and trained biologists. Tick identification was based on criteria established through

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The research for this paper was conducted at the Research Department, Maine Medical Center, Portland, Me.

Requests for reprints should be sent to Peter W. Rand, MD, Research Department, Maine Medical Center, 22 Bramhall Street, Portland, ME 04102.

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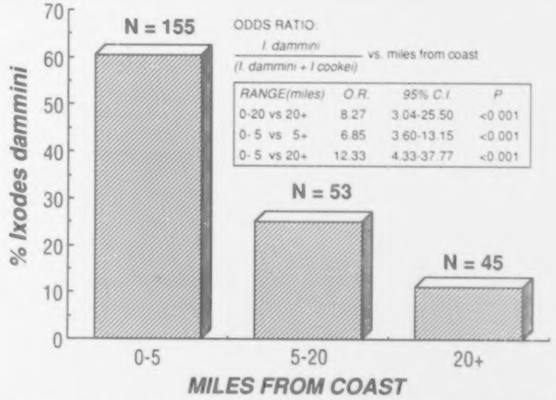
TABLE 1—Previous Canine Anti-*Borrelia* Serosurveys

Year of Study	Location	No. of Dogs	Method	% Positive	Reference
		Surveys of random asymptomatic dogs			
1984	Wisconsin	380	IFA ≥1:64	53	14
1984–1985	New Jersey	221	IFA ≥1:64	27.6 (nonendemic areas)	15
		202		42.6 (endemic areas)	
		Surveys including symptomatic dogs			
1979–1983	Southern Connecticut	210	IFA ≥1:64	28.6 (endemic areas)	16
1983, 1985	North Carolina	1002	IFA ≥1:64	3.6 (1983)	
				2.7 (1985)	17
1984–1985	Connecticut, New York	271	IFA ≥1:64	66.5 (CT)	
			ELISA OD ≥0.12	76.3 (NY)	18
1986	Northern California	345	IFA ≥1:64	26	19
1988	Texas	2400	IFA ≥1:128	5.5	20

Note. IFA = immunofluorescent assay; ELISA = enzyme-linked immunosorbent assay.



RANGE(miles)	O.R.	95% C.I.	P
0-20 vs 20+	8.27	3.04-25.50	<0.001
0- 5 vs 5+	6.85	3.60-13.15	<0.001
0- 5 vs 20+	12.33	4.33-37.77	<0.001

FIGURE 3—Percent *I. dammini* among *I. dammini/cookei* submissions from humans, cats, and dogs vs. miles from the coast.

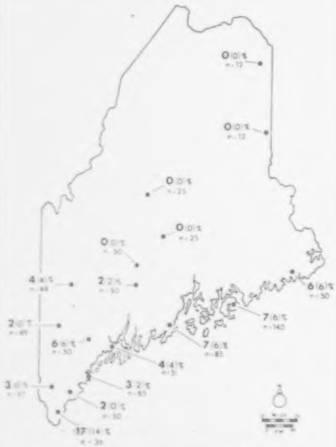


FIGURE 4—Clinic seroprevalence. Bold number is percent of all dogs in each clinic that were seropositive. Number in parentheses is percent of untraveled dogs that were seropositive. n = total number of dogs studied in each clinic.

standard keys,[8–12] verified as necessary by the Smithsonian Institution, Washington, DC.

Sixteen high-volume veterinary clinics throughout Maine (Figure 1) were invited to collaborate on the basis of their proximity to recognized endemic regions (three southernmost clinics), to areas — primarily coastal — where specimens of *I. dammini* had been collected (five clinics from Cape Elizabeth to Machias), and to inland and northern regions with little or no history of the disease or its vector (eight clinics, one serving two towns). Each clinic was provided with sampling materials and questionnaires for each pet owner. Veterinarians were asked to provide an additional 1 ml of serum from 50 or more dogs tested for heartworm microfilaria between April 1 and June 1, 1989. The only selection criteria imposed were that dogs be over 6 months old, lifetime residents of Maine, and allowed, at least sometimes, to run in fields or woods. Frozen sera were shipped to Tufts University Veterinary Diagnostic Laboratory for analysis of antibodies to *B. burgdorferi* by enzyme-linked immunosorbent assay.[13] Questionnaire items included the dog's breed, age, sex, principal function, degree of confinement, and history of travel to endemic areas. Owners were also asked if their pets had previously been lame or treated for Lyme disease.

Results

Tick Collection

Of 1605 ticks examined in 1989, 585 were *I. dammini*. Of these, 218 were submitted and 182 were collected from 789 live-trapped small mammals, 78 from mist-netted birds, 39 from flagging, and 68 from 30 of 227 examined deer. Towns from which *I. dammini* and other tick species were received are shown in Figure 2. There is notable clustering along the coast: all but 5 of 53 towns producing *I. dammini* are within 20 miles of the tidal water.

To demonstrate that the high numbers of *I. dammini* submitted from coastal regions by the public were not the result of increased population density or vigilance in these areas, submissions of this tick

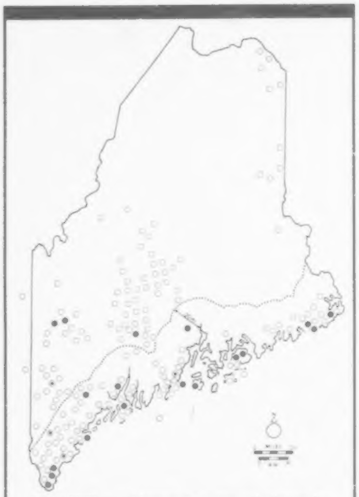
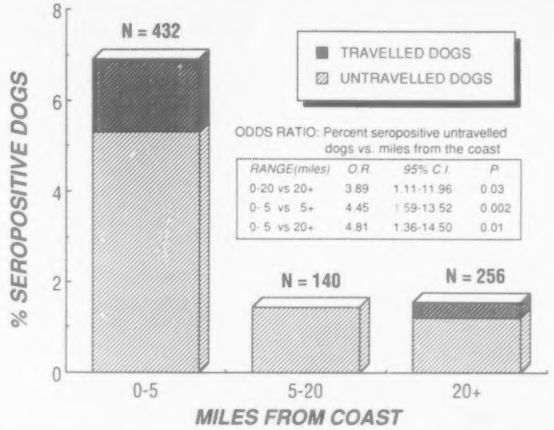


FIGURE 5—Home towns of dogs in the study. Open circles (○): no dogs positive; solid circles (●): positive untraveled dogs present; dotted circles (◎): positive traveled dog present.



ODDS RATIO: Percent seropositive untravelled dogs vs. miles from the coast

RANGE(miles)	*O.R.*	*95% C.I.*	*P.*
0-20 vs 20+	3.89	1.11-11.96	0.03
0- 5 vs 5+	4.45	1.59-13.52	0.002
0- 5 vs 20+	4.81	1.36-14.50	0.01

FIGURE 6—Percent of traveled and untraveled dogs positive for anti-*Borrelia* antibodies vs. miles from the coast.

were compared to submissions of *Ixodes cookei*, which closely resembles *I. dammini*, is widely distributed in Maine, and is found more frequently on humans than *I. dammini*. Figure 3 shows that *I. dammini* comprised 64% of the combined *I. dammini/cookei* submissions within 5 miles of the coast, but only 10% of the submissions beyond 20 miles.

Canine Serosurvey

Sera from 828 dogs were received from 149 towns representing all counties. Thirty-six samples were positive, 28 of which were from untraveled dogs. Clinics with seropositive dogs were limited to primarily coastal areas in the southern half of the state (Figure 4). Highest seroprevalence (17%) occurred in dogs from the southernmost clinic. Positive untraveled dogs were also found in two inland communities, both on rivers but in nonendemic counties. No positivity was found among 102 dogs in the northern half of the state.

Figure 5 demonstrates that, with the exception of four towns, each of which had one positive dog, the home towns of positive dogs were all within a line arbitrarily drawn 20 miles (32.2 km) from tidal saltwater. Thirty-two of all 36 positive dogs (89%) and 25 of 28 traveled dogs (89%) were within this coastal zone; in fact, the great majority of both untraveled and traveled positive dogs were within the first 5 miles (8 km) of the coast, as shown in Figure 6.

Among the positive dogs there were more retrievers than in the negative group (55% vs 40%), median age was higher (6.3 vs 4.7 years), and a greater percentage of dogs were allowed to run free (61% vs 54%). The percentage of asymptomatic animals in each group was not significantly different.

Discussion

Comparison with previous studies (Table 1) indicates that even the highest level of canine seropositivity in Maine is less than that in other endemic states. While this may reflect relatively lower tick densities and infection rates, the presence of canine seropositivity so geographically coincident with the distribution of *I. dammini* cannot be overlooked as an indication of potential risk to humans in Maine or in other emergent areas. Subsequent to this study, eight additional cases of locally acquired human Lyme disease have been confirmed in Maine, all from coastal counties. Thus, both canine serology and vector tick distribution surveys have identified areas of human risk in advance of substantiation of that risk through passive human surveillance.

Canine anti-*Borrelia* serosurveys are simple, efficient, and economical. Coordination of sampling with annual heartworm testing is particularly advantageous because it provides a large, readily available population of dogs with known histories and a standard sampling time (April and May)—an important consideration given the marked seasonal variations in canine seropositivity that have been observed.[14, 18] The introduction of a canine Lyme disease vaccine in June 1990, however, may impair the future usefulness of dog serosurveys since dogs who receive it must be excluded from serosurveys. Anything beyond minimal vaccine use in an emergent area might introduce enough selection bias to limit the epidemiologic relevance of canine serosurvey data. ☐

Acknowledgments

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NCHS to Publish New Data on Persons With and Without Health Care Coverage

The new study is entitled "Characteristics of Persons With and Without Health Care Coverage: United States, 1989" (*Advance Data* No. 201, National Center for Health Statistics, Centers for Disease Control, PHS, DHHS).

Highlights

An estimated 33.9 million Americans were without health care coverage in 1989. Health care coverage includes private health insurance, Medicare, public assistance (primarily Medicaid), and military or Veteran's Administration health benefits. Data were collected in the 1989 National Health Interview Survey, a nationwide household interview survey of the Nation's civilian, noninstitutionalized population.

People who lacked health care coverage included almost 40% of unemployed workers 18 years old or older (38.3 percent) and almost a third of people living below the poverty level (32.5 percent).

About 76.1% of the population were covered by private health insurance; 12.6% through Medicare; 6.2% by public assistance health care programs; and 2.6% were covered by military and/or Veterans' Administration health care programs.

Those without health care coverage increased from 13% in 1984 to 13.9% of the population in 1989. Although the proportion of people without coverage is greater for those below the poverty level, the increase in lack of health care coverage was greater for those above the poverty level during this period.

Situational Factors Associated with AIDS Risk Behavior Lapses and Coping Strategies Used by Gay Men Who Successfully Avoid Lapses

Jeffrey A. Kelly, PhD, Seth C. Kalichman, PhD, Michael R. Kauth, MA, Hilda G. Kilgore, MPH, Harold V. Hood, BA, Peter E. Campos, PhD, Stephen M. Rao, BA, Ted L. Brasfield, and Janet S. St. Lawrence, PhD

ABSTRACT

While most gay men have reduced behavior practices at high risk for HIV infection, there is growing evidence that many also lapse to unsafe sex. This study examined situational factors related to risk behavior lapses as well as coping strategies used by men who successfully resist lapse urges.

A convenience sample of 470 men patronizing gay bars or attending social organization meetings in four cities was surveyed. Forty-five percent of men were classified as "lapsers" (those who had had unprotected anal intercourse in the previous 6 months) and 24% were classified as "resisters" (those who successfully resisted urges to engage in this behavior). All provided information concerning the importance of factors related to the most recent occurrence of either unsafe sex or resisting unsafe urges.

Most episodes of unsafe sex occurred outside monogamous relationships and with partners of unknown HIV serostatus, although simply inquiring about partner serostatus was relatively common. Lapsers rated affectionate feelings and wishing to please a partner as well as spontaneity of unsafe sex as the most important situational factors surrounding high-risk behavior. Resisters of unsafe sex urges reported active cognitive self-guidance, experience in safe sex, and recall of both AIDS fears and safety benefits as their most important coping strategies.

Gay men who continue high-risk behavior may be overrelying on partner reports of negative serostatus. Lapse prevention approaches tailored to situations that create increased risk vulnerability must be developed. Teaching skills already used by men who successfully resist unsafe sex urges might be one approach. (*Am J Public Health.* 1991;81:1335–1338)

Introduction

Gay men in AIDS epicenters have made substantial risk reduction behavior changes.[1-5] However, these changes are difficult for many individuals to sustain. One third of homosexual men in Chicago who successfully initiate risk reduction later resume high-risk practices[6,7] Cross-sectional studies indicate that 30% to 40% of gay men in other cities still report engaging in unprotected anal intercourse in a given 3-month period.[8-10] Relapse is a concern even among gay men enrolled in San Francisco AIDS behavioral research cohorts, presumably one of the most AIDS-sensitized samples in the United States.[11]

Homosexual men who continue to engage in high-risk sex even occasionally are in great danger of contracting HIV infection. Gay men in large cities who have unprotected receptive anal intercourse with a single partner over 1 year are three times more likely to seroconvert than men who refrain from this activity and, with five partners, are 18 times more likely to develop HIV infection.[12] While lapses are the rule rather than the exception following initiation of change in all health behavior areas,[13-15] we were aware of no other population or health area where there is such a small margin for error and where even infrequent lapses carry such potentially grave health consequences.

Risk behavior relapse among gay men is predicted by younger age, lower education level, excessive intoxicant use, frequency of past risk behavior with multiple partners, and high reported past enjoyment associated with risky practices.[8,11,16,17] The present study more closely evaluated specific situational and social context factors surrounding risk behavior lapses as well as the coping strategies used by gay men who report successfully resisting high-risk behavior urges.

Method

Subjects

Subjects were 406 men who patronized gay bars and 64 men who attended social meetings of gay organizations in the spring of 1990 in four cities: Memphis, Tenn, Tampa, Fla, Mobile, Ala, and Binghamton, NY. Approximately 75% of men who were approached agreed to answer questions anonymously. The mean age of respondents was 31.0 years (SD=8.6) and the mean education level was 14.9 years (SD=2.2); 91% of the sample was White and 9% was Black, Hispanic, or of other racial or ethnic backgrounds. The city samples did not differ significantly in demographic characteristics.

Measures

Respondents described their sexual behavior over the preceding 6 months including frequency and number of partners for unprotected anal intercourse, both receptive and insertive. Forty-two percent of men who engaged in unprotected anal intercourse were the insertive partner, 35% the receptive partner, and 23% were both insertive and receptive during their most recent behavior. Multivariate analysis of variance revealed no significant difference between those who were the in-

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Jeffrey A. Kelly and Ted L. Brasfield are with the Medical College of Wisconsin, Seth C. Kalichman is with Loyola University of Chicago, Michael R. Kauth is with the University of Mississippi, Hilda G. Kilgore and Harold V. Hood are with the Tampa AIDS Network, Peter E. Campos and Stephen M. Rao are with the State University of New York at Binghamton, and Janet S. St. Lawrence is with Jackson State University.

Requests for reprints should be sent to Jeffrey A. Kelly, PhD, Department of Psychiatry and Mental Health Sciences, Medical College of Wisconsin, 8701 Watertown Plank Road, Milwaukee, WI 53226.

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sertive vs receptive partner in importance ratings for lapse situation items, so the insertive and receptive groups were combined for all data analyses. They were also questioned about number of partners and frequency for anal intercourse using condoms and about safer sex practices without penetration such as mutual masturbation and frottage (body rubbing). In addition, all men reported whether they presently had condoms in their pocket, car, and home.

TABLE 1—Situations Associated with the Most Recent Occurrence of High-Risk Behavior among Gay Men Who Report Unsafe Sex during the Previous 6 Months

Situational Descriptor	Mean (SD) Importance Rating
This partner was someone special to me	2.7 (1.3)
I wanted to please this partner	2.2 (1.3)
Once sex began, we got caught up in the passion of the moment	2.2 (1.2)
No condoms were nearby at the time of sex	2.0 (1.3)
Using a condom seemed impersonal	1.9 (1.1)
Using a condom would have reduced sensation/pleasure	1.9 (1.2)
Stopping to use a condom would have interrupted sex	1.8 (1.1)
I had a strong urge to have anal intercourse without a condom	1.8 (1.1)
It was hard for me to refuse having anal intercourse	1.8 (1.1)
It was a long time since I had anal intercourse without a condom	1.8 (1.1)
I had a lot to drink or used drugs and wasn't clearheaded	1.7 (1.1)
Condom use implies a lack of trust	1.5 (0.9)
It was embarrassing for me to suggest using a condom	1.5 (0.9)
I felt no control over what we did	1.5 (1.0)
I am not able to change my sex practices	1.5 (0.9)
I was feeling depressed, lonely, or upset at the time	1.4 (0.9)
This partner pressured me not to use a condom	1.4 (0.9)
It was too hard to get the condom package open	1.3 (0.7)
I didn't have the energy to use a condom	1.3 (0.6)
I pressured this partner not to use a condom	1.2 (0.6)
I was too embarrassed to buy condoms	1.2 (0.7)

Note. Importance ratings made on scale from 1=not a factor to 4=a big factor.

TABLE 2—Coping Strategies Reported to Be Used by Gay Men Who Successfully Resisted Urges to Engage in Unsafe Sex during the Previous 6 Months

Coping Strategy	Mean (SD) Importance Rating
I remembered information about AIDS and safer sex	3.6 (0.9)
I thought about how important it is to be in good health	3.5 (0.9)
I was afraid of getting (or giving) the AIDS virus	3.5 (1.0)
I reminded myself to stay safe before or during sex	3.4 (1.0)
I knew how good I'd feel about myself later if I stayed safe	3.3 (1.1)
I didn't want to worry about AIDS after sex	3.3 (1.1)
I am experienced in safer sex	3.3 (1.0)
I decided ahead of time what I would or would not be willing to do	3.2 (1.0)
I actively guided our actions to stay safe	3.1 (1.0)
The scare of knowing people with AIDS motivated me to be safe	3.1 (1.2)
I told this partner we needed to practice safer sex	2.9 (1.1)
I wanted this partner to know I was health conscious	2.9 (1.2)
I kept condoms nearby	2.9 (1.2)
I told this partner I didn't want to have unprotected intercourse	2.9 (1.2)
We switched from something that was getting risky to something safer	2.5 (1.2)
Safer sex was as good as intercourse without a condom	2.5 (1.2)
Safer sex was expected by this partner	2.4 (1.3)
I didn't drink or use drugs much before sex so I could be clearheaded	2.3 (1.1)
This partner told me he didn't want to have unprotected intercourse	2.3 (1.2)
This partner told me that we needed to practice safer sex	2.2 (1.3)
I didn't want other people to find out that I might have unsafe sex	2.0 (1.2)

Note. Importance ratings made on scale from 1=not a factor to 4=a big factor.

Respondents who reported engaging in unprotected anal intercourse in the previous 6 months ("lapsers," n=209, 45% of the sample) described, for the most recent occurrence of high-risk behavior, exclusivity of relationship with the partner, knowledge of own and partner's HIV serostatus, difficulty in maintaining safer sex practices, and percentage of anal intercourse occasions when condoms were used in the previous year. These men then rated the importance of 21 situational factors associated with the most recent occasion of high-risk behavior. As Table 1 shows, these items assessed environmental and setting factors, feelings toward the partner, coercion, physiological and emotional state, management of sudden urges, and self-control.

A second set of men were those who reported successfully resisting temptation to engage in unprotected anal intercourse in the previous 6 months ("resisters," n=116, 24% of the sample). (Men who reported no occurrence of unprotected anal intercourse *and* no temptation to engage in the practice were not studied here.) This group of men rated the importance of 21 coping strategies that may have been used on the most recent occasion when high-risk urges were resisted. Table 2 lists these items, which include risk reduction self-prompts; environmental modifications; active use of alternative behaviors; cognitive statements emphasizing benefits, fears, or beliefs of personal efficacy; and social expectations to avoid risk.

Results

As Table 3 shows, men reporting high-risk behavior were younger than those who reported resisting unsafe sex urges, had less education, and had had more male sexual partners in the previous 6 months. Groups were comparable in reporting having condoms in their pocket, car, or home and reported similar difficulty levels for remaining safe during sex.

Table 1 presents the mean importance rating for each situational descriptor associated with high-risk behavior by men who reported lapses, ranked from most to least important. Factors most critical to lapses were positive emotional feelings for the partner, wishing to please the partner, not having condoms nearby at the time of sex, becoming "caught up in passion of the moment," and negative connotations about condom use. As Table 2 shows, men who resisted high-risk urges reported the following as their most important coping strategies: cognitive reminders to re-

TABLE 3—Characteristics of Men Who Lapsed to Unsafe Sex and Men Who Successfully Resisted Relapse Urges during the Previous 6 Months

Characteristic	Lapsers (n=209)	Resisters (n=116)	95% CI[a]
Mean age in years (SD)	29.4 (7.6)	31.9 (8.4)	29.5–31.3
Mean educational level in years (SD)	14.6 (2.1)	15.6 (2.0)	14.7–15.2
Mean number of different male sexual partners in previous 6 months (SD)	4.5 (7.6)	3.4 (4.2)	3.4–4.8
Percentage of men reporting having condoms in their			
Pocket or wallet	29	27	—
Car	42	37	—
Home	84	93	—
Mean personal difficulty rating for remaining safe during sex (SD)[b]	2.6 (1.2)	2.5 (1.1)	2.3–2.8
Mean percentage of intercourse occasions when condoms were used during previous year	44	—	—

[a] CI = confidence interval.

[b] Difficulty ratings made on a 5-point Likert scale from 1=very easy to stay safe to 5=very difficult to stay safe during sex.

TABLE 4—Relationship between Knowledge of Own HIV Serostatus and Knowledge of Sexual Partner's Serostatus among Men Reporting Occurrence of Unprotected Anal Intercourse

	Knowledge of Partner's HIV Serostatus			
Knowledge of Own HIV Serostatus	Didn't Know Partner Serostatus (n=97)	Partner Knew Was HIV Positive (n=5)	Partner Said He Was HIV Negative (n=61)	Saw Evidence Partner Was HIV Negative (n=35)
Not tested (n=55)	44 (22%)	1 (0.5%)	9 (4.5%)	1 (0.5%)
HIV positive (n=10)	7 (4%)	1 (0.5%)	2 (1%)	0 (0%)
HIV negative (n=133)	46 (23%)	3 (1.5%)	50 (25%)	34 (17%)

Note. Chi-square=40.1, df=6, $P < .0001$ for the association of knowledge of own serostatus and knowledge of partner's serostatus.

main safe during sex, self-statements regarding fear and avoiding worry, identifying the benefits of safety and good health, and experience and competence in actively guiding safer sex.

To examine the context of high-risk episodes, relationship status and serostatus knowledge between partners were also explored. Most high-risk activity (57%) took place between nonexclusive and casual partners; 43% occurred with partners described as exclusive, although the mean duration of the relationships was only 12.5 months. Table 4 shows the distribution for high-risk behavior in relation to knowledge of one's own and one's partner's HIV serostatus. Nearly half of the respondents who engaged in high-risk behavior had no knowledge of their partner's HIV serostatus; an additional 31% reported that the partner simply said he was HIV negative. In less than 18% of instances did respondents report first-hand knowledge of partner's HIV-negative status and, even when high-risk behavior occurred with a partner described as exclusive, two thirds of these men said they did not know both their own and their partner's serostatus. In general, knowledge of one's own negative status was related to a greater likelihood of inquiring about a partner's HIV status.

Discussion

Nearly one half of respondents in this sample had engaged in unprotected anal intercourse in the previous 6 months, usually with nonexclusive partners of unknown HIV serostatus. Even men who engaged in high-risk behavior appeared cognizant of risk: well over 80% said they had condoms at home and lapsers used condoms almost half the time. In addition to the association of risk behavior with affectionate feelings toward the partner and spontaneity of risky sexual practices, a large proportion of lapsers accepted partner claims of negative HIV status or their own past negative result. Among persons who continue to engage in high-risk behavior, past negative HIV results of self or partner, even if accurate, carry little current protective value. Approaches are needed that better prepare gay men to anticipate and handle lapse temptations, counter misconceptions that may perpetuate vulnerability to lapses, and stress the importance of maintaining consistent behavior change. These results indicate that resistance to lapses is related to active use of cognitive self-guidance, recall of both AIDS fears and safety benefits, and experience in safer sex. Although the current study is limited by its use of a convenience sample of unknown generalizability, respondent self-reports of behavior, and analysis only of "within-group" patterns, it highlights the need for further study of ways to prevent risk lapses in populations vulnerable to HIV infection. ☐

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Call for Papers on Issues and Innovations in Financing and Reimbursement of Health Promotion and Education Programs

Health Education Quarterly, the official journal of the Society for Public Health Education, announces a call for original manuscripts for a theme issue on "Issues and Innovations in Financing and Reimbursement of Health Promotion and Education Programs." Manuscripts on a wide variety of topics addressing this theme are solicited. All manuscripts will be peer-reviewed by a guest editorial board. Manuscripts may address the following:

- Innovative financing strategies for patient education and health promotion
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- Innovative linkages with Medicaid and Medicare
- Federal and state health promotion policy related to funding, financing, and reimbursement
- Payment strategies in various medical care settings
- Consumer payment for services
- Effects of reimbursement on the demand for and supply of patient education services
- Implication for financing of patient education under proposed changes in outpatient payment
- Other

Review papers, case studies, research studies, policy analysis, and other types of papers will be accepted. Papers covering financing, reimbursement, payment, and funding of health promotion and education in all settings, including worksite health promotion, patient education, community health promotion, and other settings are encouraged.

Papers are due December 15, 1991. Please submit a letter of intent, if possible, by July 15, 1991 (papers submitted without a letter of intent will still be accepted but letters of intent will help in the review process).

Please send the letter of intent and manuscript to Randy Schwartz, MSPH, Director, Division of Health Promotion and Education, Bureau of Health, Department of Human Services, State House Station 11, Augusta, ME 04333.